AccessLab

Regional Labour Market Adjustments in the Accession Candidate Countries

Workpackage No. 4

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Human Capital, Spatial Mobility, and Lock-In – The Experience of Candidate Countries

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Deliverable No. 6 to 10

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EI - Hungarian Academy of Sciences - Institute of Economics
IZA – Institut für die Zukunft der Arbeit
UCL/SSEES - University College London - School of Slavonic and East European Studies
ZEI - Centre for European Integration Studies
ZEW - Centre for European Economic Research

AccessLab

The 5th framework programme research project ACCESSLAB researches the capability of candidate countries' regions to deal with asymmetric shocks. Its goal is to provide analysts and policy makers with research results relevant to the process of enlargement. The project takes a broad and comparative view of labour market adjustments to address these issues. It examines the topic from both a macroeconomic and microeconomic viewpoint. It considers different adjustment mechanisms in depth and compares results with the European Union. It draws on a) the experiences in transition countries in the last decade, b) the experience of German integration and c) the experiences of border regions to gain insights on the likely regional labour market effects of accession of the candidate countries.

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Human Capital, Spatial Mobility and Lock in – The Experience of the Candidate Countries: Executive Summary and Policy Conclusions

Janos Köllö and Peter Huber

Introduction

Compared to EU-15 the countries of Central and Eastern Europe exhibit a high degree of ethnic diversity, and their societies are deeply divided by regional frontiers. Inequalities by human capital endowments have also been rising and, in several CEEs, reached levels comparable to the US rather than continental Europe. Furthermore, the candidate countries as well as the new member states have experienced substantial changes in labour market institutions in the course of transition from the planned to market economies in the last one and a half decades. Given these developments and particularities it is no surprise that a substantial body of literature has developed which analysis the effects of changes in labour market institutions on labour market performance in new member states and candidate countries. In particular the connection between active and passive labour market policies and labour supply as well as changes to the returns to schooling have been intensively researched. Furthermore, differences in labour market institutions have been analysed in detail (see. Workpackage 1 for a survey and the contribution by Ederveen and Thissen (2004) in workpackage 4). In general this literature suggests that:

1. Labour market institutions do not differ dramatically from those of many OECD countries in the new member states and candidate countries.
2. Wage inequality at the regional as well as the individual level has risen substantially in all new member states and candidate countries. This was to a substantial part caused by an increase to the returns to schooling.
3. The effect of unemployment and social benefits on labour supply is ambiguous in most candidate countries and new member states and findings depend strongly on methodological choices and episodes analysed in research.
4. Active labour market policies have varied substantially in their success in integrating unemployed into employment.

Other aspects of labour supply behaviour in candidate countries and new member states by contrast are much less explored. In particular issues related to the potential discrimination of members of individual ethnic minorities have been much less explored (see Kroncke and Smith, 1999 for an exception). Similarly the analysis of migration behaviour in these countries has only recently received more attention (see workpackage 3 of the AccessLab project) and evidence on commuting is scant for most of the new member states and candidate countries (see Kertesi (2001) and Hazans (2003) for exceptions)

The Contents of this Report

The objectives of workpackage 4 of the AccessLab project were to extend on the literature on labour supply in the candidate countries and to contribute to filling some of the gaps by determining the reasons for regional "lock-in" in candidate countries, providing insights into how different demographic groups are affected by (and react to) these shocks and assessing migration and commuting behaviour in the labour markets in the candidate countries and new member states.

In particular a number of contributions in the workpackage (chapters 1 and 2 as well as chapter 7) are devoted to identifying the impact of labour market institutions and systemic changes on different aspects of labour market performance. Kertesi and Köllö (in chapter 1) focus on the effects of a particularly spectacular case of increases in minimum wages in Hungary. Andren, Earle and Sapatoru (chapter 2) focus on the effects of systemic reforms on the returns to schooling in Romania and Hazans (chapter 7) isolates the effects of changes in labour market policy in the Baltic countries on the labour supply decision.

Furthermore the contributions of Workpackage 4 study the emerging and/or already existent social frontiers in the new member states, by using micro data from different countries and time periods. A central concern in this respect is the role of ethnic minority members in the labour market. Kertesi (in chapter 8) presents a detailed study of the labour market situation of the Roma in Hungary and Smith (in chapters 9 and 4) as well as Hazans (chapter 7) consider the labour market situation of ethnic Russians in terms of wages and employment prospects in the Baltic countries. The analysis of the impacts of policies on different demographic groups, however, is also discussed from a perspective in many of the contributions. Smith (in chapter 3) and Andren Earle and Sapatoru (in chapter 2) highlights the role of increasing returns to education and experience in determining wages, Hazans (in chapter 7) stresses the particular role of the elder in explaining labour supply reductions in the Baltics, while the contributions on commuting and the willingness to migrate by Bartusz (in chapter 5) and Fidrmuc and Huber (in chapter 6) stress the role of gender and education in shaping individual attitudes to mobility in the new member states and candidate countries.

Finally, a number of contributions to workpackage 4 extend on the previous analysis of regional mobility in the new member states and candidate countries provided in workpackage 3. While Huber (in chapter 4) presents a comparison of place to place migration rates and thus extends on the analysis provided in workpackage 3 by Fidrmuc (2003), Bartusz (in chapter 5) analysis the commuting decision of unemployed job finders in Hungary, thus filling an important gap in the literature on labour market adjustment in candidate countries and new member states, and Fidrmuc and Huber (in chapter 6) provide evidence on the individual and regional determinants of the willingness to migrate. Finally, Bruecker and Truebswetter (in chapter 10) shift the focus somewhat by analysing the impact of brain-drain on the East-German labour market after unification, thus providing important insights on the potential effects of such brain drain on the new member states after enlargement.

Results

Given the nature of the analysis, the data requirements, and the wide focus of topics covered, the workpackage did not aim at broad coverage and/or cross-country comparison. The papers rather tried to benefit from the richness of individual and firm-level data providing insight to the issues analysed extensively in Workpackages 1 to 3. Given this focus the results may be summarized as follows:

1. Increasing minimum wages does not seem to contribute to reduced unemployment. Although total employment seems to have been only marginally affected, Kertesi and Köllö (in chapter 1) suggest that minimum wage increases in Hungary significantly increased labour costs, reduced employment in the small firm sector, and adversely influenced the job retention and job finding probabilities of low-wage workers. Furthermore, higher minimum wages also seem to be an inefficient instrument in reducing regional disparities. Depressed regions were equally or more severely hit, suggesting that the demand-side reactions dominated everywhere. While this suggests that higher minimum wages are unlikely to yield substantial improvements in terms of unemployment, they may contribute to higher labour force participation in some cases. Hazans (in chapter 7) finds some evidence that increasing minimum wages led to higher participation and reduced the share of discouraged workers in the workforce in Lithuania. In Estonia by contrast increased participation is only found for teenagers and young males.
2. Discrimination on ethnic grounds hampers regional labour market adjustment in the candidate countries and may be considered an important element causing regional "lock-in". The region’s division by

ethnicity, language and religion manifested itself in several ways since 1989 even including tragic inter-ethnic hostilities. The EU accession countries experienced less of the open conflicts but several of them have to cope with severe inequalities related to ethnicity. Hungary, Slovakia and the Czech Republic have sizeable Roma minorities living in underclass conditions and facing several times higher unemployment rates than do the non-Roma. The Baltic States through their large Russian minority are also challenged by a minority problem that is unparalleled in its scope and nature within the former EU. Roma are the largest low status ethnic minority of Central Europe and the Balkans, and their deprivation represents one of the region's most severe social problems. Using survey data the Kertesi (in chapter 8) in his in-depth analysis of the exclusion from the labour market of Hungary's sizeable Roma community (accounting for about 6 per cent of the country's population) suggests that under socialism (1984) 75 per cent of the Roma adults were steadily employed in large industrial organisations rather than traditional Gypsy occupations. By 1994 their employment ratio fell to 35 per cent and remained at that level until recently.[1] The study demonstrates that those staying in employment also have shorter job spells. About half of the employment gap can be attributed to lower education of the Roma, and their regional affiliation adds a further compositional effect. Industry-specific shocks do not explain the residual gap given that the Roma were not over-represented in industries severely exposed to the transition shock. Both the time path and the regional patterns of Roma unemployment suggest, however, that they were 'crowded out' by majority workers on a massive scale. Roma employment started to decline prior to 1989 as Hungary introduced a series of market institutions and hardened the enterprises' budget constraints. The bias against Roma workers also appears in their *relative* employment rates across regions. The employment gap between the Roma and the non-Roma sharply increases with the local unemployment rate - an observation that is hard to reconcile with non-discriminatory practices.

Segregation in education also seems to play a role in transmitting the disadvantageous position of ethnic Russians in Estonia and, less clearly, Lithuania. For the Baltic countries Smith (in chapter 9) identifies substantial ethnic earnings wage gaps in Estonia and Latvia, and lower returns to human capital for members of the Russian minority in Estonia and Lithuania. In Estonia the bulk of the earnings gap is

[1] The 2001 Census, for instance, suggested that the Roma population's employment ratio fell short of 1/3 of the country's aggregate employment ratio.

attributable to differential returns to human capital, which is partly explained by the lower quality of Russian language education. The case is different in Romania where Andren, Earle and Sapatoru (in chapter 2) find no statistically significant gaps in returns to human capital comparing ethnic Romanians, Hungarian and Germans. In particular their findings on the Romanian labour market refute the hypothesis that minorities' higher potential to migrate leads to higher wages and/or higher returns to human capital.

3. Aside from discrimination on ethnic grounds marked differentiations exist for labour market outcomes among different groups, which suggests substantial room for micro-oriented labour market policies. In particular returns to education increased dramatically during transition, which caused wage inequality to increase substantially. Furthermore lowly qualified workers are the main group with the largest difficulties in adjusting to labour market shocks. Andren, Earle and Sapatoru (in chapter 2) find that in Romania returns to schooling increased from 4% in socialist times to 8.5% in 2000 and Smith (in chapter 3) finds similar stylised facts concerning household income in the Baltic countries. Addressing a number of alternative hypotheses concerning the increase in returns to schooling Andren, Earle and Sapatoru (in chapter 2) conclude that the high productivity of school-based skills (pre- and post-transition alike) in restructuring and entrepreneurial activities played key role in the doubling of returns to education.

   Low skilled workers are, however, also found to be disadvantaged in a number of further respects relevant to their labour market adjustment. They are likely to have the lowest willingness to migrate (see Fidrmuc and Huber in chapter 6) and have lower chances of moving between labour market states (see: Hazans in chapter 7). While this finding is in accordance with much of the literature on labour market adjustment in the old EU member states it does suggest that issues of the education of the workforce are an important aspect in facilitating regional labour market adjustment.

4. Some evidence also suggests that gender differences in labour market outcomes may be further reason for regional lock in labour markets. In particular Bartusz (in chapter 5) finds that commuting distances of Hungarian job finders are lower for females than for males. Furthermore Hazans (in chapter 7) reports that women in particular in the lower qualification strata are more the most likely demographic group to be discouraged unemployed.

5. Lacking regional mobility in the candidate countries is an important element in explaining the persistence of regional disparities in the new member states and candidate countries. Evidence on the responsiveness of spatial mobility to unemployment and wage differentials suggests a low propensity to migrate. Migration flows have been declining during the transition, even as regional disparities have been rising. Estimating place to place migration Huber (in chapter 4) finds that migration is less reactive to regional disparities in accession countries than in EU-15 states. If reaction to disparities were similar gross migration should increase by 10 to 50 per cent and net migration by a factor of between 2 and 10. A number of reasons can be given for these findings. First, as already found in the results of workpackage three owner occupied housing, high mobility costs and vacancy chain effects have kept migration and commuting at relatively low levels in the CEEs.
6. As shown in the analysis of the willingness to migrate in the Czech Republic by Fidrmuc and Huber (chapter 6) individual willingness to migrate depends more strongly on personal characteristics rather than on the regional labour market situation. In particular females and less qualified persons have a lower willingness to migrate. Fidrmuc and Huber (chapter 6) also find that ownership of either own housing and or weekend houses seems to limit willingness to migrate. While these results require some corroboration before jumping to strong policy conclusions, this suggests capital and housing market inefficiencies seem to play some role in explaining low migration in the candidate countries and that improved human capital will increase the adaptability of the workforce.
7. Commuting, a potentially viable alternative to migration is constrained by high transport costs relative to wages and bottlenecks in public transport connections. The contribution by Bartusz (in chapter 5) finds that travel to work costs severely constrain the commuting distance of unemployed workers in Hungary. Long-distance commuting seems conditional on employers' contribution to travel to work costs with only 15 per cent of the commuters self-financing their travels. Estimating a model of commuting decisions we find that travel to work costs limit the distance of self-financed commuting to 20 km with women and 50 km with men. These findings are similar to those of earlier research by Kertesi (2001), who found that commuting costs tend to lock low-wage workers into high-unemployment villages while high-educated residents are able to access urban labour markets, and also Hazans (2003) on the Baltic countries, which suggests that inter-community commuting is also low in the Baltic countries.

8. The consequences of the selectivity of migration with respect to education may have implications for the sending regions. This is the result of Brueckers and Truebswetters study (in chapter 10) on the effects of "brain drain" after German Unification on regional development. In contrast to a number of studies investigating "brain drain", they find some indication of a negative effect of "brain drain" for the immobile residents of a region. East German workers can realise higher wage growth if there is a high share of highly qualified in the same district and a lower emigration of highly qualified out of this district. For immobile workers in sending regions it would thus be preferable to restrain qualified worker from migrating.

Policy Conclusions

Clearly the rich results of the reports included in workpackage 4 indicate a number of important policy implications for the countries analysed. While it seems difficult to generalise the results of individual studies to a set of countries that differ substantially in their institutional, economic and social situation as the candidate countries we would argue that the most important policy lessons to be learned from this workpackage are, that:

- Fighting the disincentives to individual adjustment that inevitably develop in low-wage environments requires careful policies addressing demand-side deficiencies and transaction costs, rather than aggregate level policy intervention aimed at labour supply. This is evidenced by the study on the natural experiment of doubling the minimum wage in Hungary 2001-2002, which was a straightforward attempt to break low equilibrium by widening the gap between wages and benefits. In extension of this result one could expect that other more macro oriented policies directed at increasing search incentives for the unemployed (such as reductions in unemployment benefit entitlements) are also unlikely to contribute to reducing high unemployment in particular in regions with low labour demand. A suggestion that is also stressed in workpackage 5 in the contribution by Ederveen and Thissen (2004) who find that an approach focusing on labour demand deficiencies, combating skill mismatch and improving policy implementation are likely to be the most efficient in reducing regional labour market problems.
- Furthermore, some scepticism concerning the potential of such aggregate policies to reduce regional disparities seems to be warranted. At least in the Hungarian minimum wage experiment depressed regions were equally or more severely hit by the hike despite the fact that some positive supply-side effects, as predicted in several theoretical models of the minimum wage, are more likely to develop under conditions

characteristic of such provinces. (Workers have higher probability of receiving unemployment benefits; the benefits replace a larger fraction of their lost earnings; they have better than average access to informal second jobs, are more severely constrained by fixed costs like travel-to-work expenses whereas monopsonies are also more likely to occur.) The evidence thus suggests that even in these regions the expected positive supply-side responses were more than offset by the elementary cost effect of a move to a higher minimum wage. We thus conclude that as long as the equilibrating mechanisms of the labour market work sluggishly the depressed regions face a high risk of slipping to a low equilibrium state characterised by low participation and wages, and massive reliance on social welfare. Thus we would also argue that a policy addressing the issues of regional demand deficiencies and investments into an improved implementation of regional policy are more likely to contribute to regional equality.

- Minority issues are and will be a major issue in the policy debate on social cohesion in the new member states as well as in the candidate countries for some time to come. The findings, in sum, call for action in educational and regional policies as well as in the enforcement of anti-discrimination laws. The degree and nature of social exclusion demonstrated in the individual papers warns that the re-integration of the Roma (in the CEEs as well as the Balkans) should be given high priority in an EU committed to social cohesion. Fighting school segregation seems particularly important in order to block the inter-generational transmission of deprivation.
- As in the old EU member states education policy and strategies to implement life long learning seem to be a key element in facilitating the adaptability of the workforce in new member states and candidate countries. While in this respect both candidate countries and new member states do not differ much from the old EU-member states, we would argue that the priority given to designing efficient strategies of increasing the human capital stock in these countries (and in particular in backward regions) should even be higher in the new member states and candidate countries than in the old member states, because the dramatic increases in returns to education and the low mobility of less skilled workers, suggest substantial skill mismatch in the regional labour markets in new member states and candidate countries.

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MINIMUM WAGE AND EMPLOYMENT – HUNGARY'S EXPERIMENT *

GÁBOR KERTESI - JÁNOS KÖLLŐ

The Hungarian government's decision to double the minimum wage within a year provides a unique opportunity of testing how employment reacts to an exogenous shock to wages. The short-run effect of a fifty-seven per cent rise in the minimum wage in January 2001 is analyzed using matched employer-employee data and panels of labor market flows. The hike significantly increased labor costs, reduced employment in the small firm sector, and adversely influenced the job retention and job finding probabilities of low-wage workers. Depressed regions were equally or more severely hit, suggesting that the demand-side reactions dominated everywhere. The results yield support to the competitive framework in predicting minimum wage effects albeit, consistent with findings from elsewhere, they indicate a minor change in aggregate employment.

I. INTRODUCTION

In January 2001 the Hungarian government increased the minimum wage from monthly Forint (Ft) 25,500 to Ft 40,000. One year later the wage floor rose further to Ft 50,000. There were few examples for adjustments of this magnitude in recent economic history, with Puerto Rico and Indonesia being well-documented exceptions. (See Castillo-Freeman and Freeman [1992] on the former and Rama [2000] and Alatas and Cameron [2003] on the latter).

This paper looks at the short-run employment effect of the first hike. The theoretical predictions are ambiguous and have been subject to a reviving debate since the publication of David Card and Alain Krueger's 'Myth and Measurement' in 1995. The 'new economics of the minimum wage' predicts positive employment effect in a variety of market structures including monopsony (Ehrenberg and Smith [2000] and other textbooks), distortions under monopsonistic competition (Bashkar et al. [2002]), efficiency wage setting (Rebitzer and Taylor [1995]), search frictions (Ahn and Arcidiacono [2003]) and dual wage determination (Cahuc at al. [2001]). Furthermore, it was long ago demonstrated by Mincer [1976] that equilibrium employment can rise and unemployment fall depending on how the elasticities of demand and supply relate to each other and the labor turnover rate.

The decision to radically adjust the minimum wage in Hungary was undoubtedly motivated by some of the unorthodox considerations. The motives were presented in popular form (the hike will 'restore the prestige of work', combat the misuse of benefits, 'whiten the black economy', and so on) but the political slogans actually drafted some key arguments of

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the new theories. It was argued that by widening the gap between wages and benefits the government can create proper incentives for paid employment, encourage job search, promote competition for job openings and stimulate work effort. Any negative effect resulting from higher wages could thus be offset by the returns to better incentives and falling transaction costs .[1]

Indeed, a series of empirical papers including Card [1992a,b], Katz and Krueger [1992], Card and Krueger [1994, 1995], Machin and Manning [1994] and Dolado et al. [1996] observed close to zero or positive change in employment after minimum wage hikes in the US and Europe. In Costa Rica El Hamidi and Terrell [1997] found the impact of hikes to be positive at low levels of the minimum wage but negative in higher ranges of the industrial minimum wage-average wage ratios. The time series evidence from the 1990s also suggested significantly weaker negative effect than those found earlier (Brown [1999]). However, a whole array of papers continued to identify significant negative impact including Kim and Taylor [1995], Deere et al. [1995] and Neumark and Wascher [1994, 2002] in the US, Abowd et al. [1999] in a US-France comparison, Bell [1997] and Maloney and Mendez [2003] in Colombia, Castillo-Freeman and Freeman [1991] in Puerto Rico, Pereira [1999] in Portugal; Rama [2000] and Alatas and Cameron [2003] in Indonesia. The effects found in these studies are often small and restricted to certain segments of the market like teenagers and small firms but they definitely lend support to the orthodox predictions.

Given the ambiguity of the theoretical predictions and controversy over the 'stylized facts' the analysis of an extraordinary rise in the minimum wage may contribute to the ongoing debate. This, we believe, remains true in view of the fact that it is difficult to identify the effect of changes in a *single* national minimum wage. The difference-in-difference approach, which relates differences in the outcomes to differences in the treatment of otherwise identical actors, is clearly not applicable in this case. When a single minimum wage is adjusted all the variation in exposure is explained by variation in the pre-hike wages of firms or individuals - supply and demand shocks affecting low-wage and high-wage workers in a different way can thus establish spurious correlation between exposure and employment outcomes. Ignoring the case of national minimum wage legislation is one option for

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[1] The stereotype of general support on the political left and opposition on the right does not apply in this case. The hikes were decided by a right-wing government explicitly committed to increasing the welfare of the middle class and promoting the competitiveness of domestic businesses including exporters - an unusual candidate for

researchers. Finding a second best strategy of identification is another and this paper makes attempts at the latter.

Section II introduces Hungary's minimum wage hike in more detail including descriptive statistics on the subsequent changes of wages and employment. Section III analyses employment in small firms - a sector severely affected by the minimum wage hike and providing exceptionally rich matched employer-employee data for 2000 and 2001.[2] Less reliable results on medium-sized and large enterprises are also presented. The conclusion that the effect of the minimum wage hike was negative and heavily concentrated on the low-wage workers of low-wage firms is further tested in Section IV, which analyze the job retention and job finding probabilities of low-wage workers.[3] The results confirm that raising the minimum wage came at the cost of low-wage job opportunities. Since the analytical parts use different types of data and models the issues of identification and other methodological problems will be discussed in the relevant sections. The data sources are introduced in the Appendix.

II. THE MINIMUM WAGE SHOCK – MAGNITUDE, COMPLIANCE, AND IMMEDIATE AFTERMATHS

The Hungarian minimum wage, introduced in 1989 by the country's last communist government, relates to monthly pre-tax 'base wages' net of overtime pay, shift pay and bonuses, is legally binding, and covers all employment contracts. For part-timers accounting for 3.5 per cent of paid employment the wage floor is proportionally lower. In 1990-1998 adjustments were negotiated annually by a national-level tripartite council while in 1998-2002 the minimum wage was set unilaterally by the government.

In May 2000 when the plan of a radical adjustment was first announced the minimum wage-average wage ratio stood at 29 per cent, a level deep below the European average but only marginally lower than Spain's, the laggard within the EU. Despite its low Kaitz-index Hungary's minimum wage *was* effective. The fraction of workers paid 95-105 per cent of the

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aggressive minimum wage policies. The largest trade union federation of socialist orientation worried about the potentially adverse employment effects openly opposed the first hike.

[2] An estimated 70 per cent of Hungarian minimum wage workers are employed in small firms with less than 50 employees.

[3] Looking at flows between labor market states is justified by a high concentration of minimum wage workers in jobs with short tenures. According to Labor Force Survey data from April-June 2001 about 20-25 per cent of the minimum wage workers had tenures shorter than one year, nearly 40 per cent worked less than 2 years, and 60 per cent had less than 5 years with the firm - while only 4.4 per cent spent less than 5 years on the labor market. Only 20 per cent were younger than 25 and a mere 2 per cent teenager. Minimum wage effects are more likely to be observed at the margin between employment and unemployment than by looking at youth employment, as most studies do in western market economies.

minimum amounted to 5 per cent - a ratio similar to those reported for Austria, Belgium, the Netherlands, Denmark, and the US by Dolado et al. [1996].[4]

The Ft 40,000 minimum wage cut deep into the wage distribution: an estimated 21 per cent of the employees ought to have received a wage lift of 28 per cent on average in order to comply with the regulations. *Figure I* based on matched wage observations from the May 2000 and 2001 waves of the Wage Survey (WS) yields an approximation of how actual base wages changed along the base-period wage distribution.

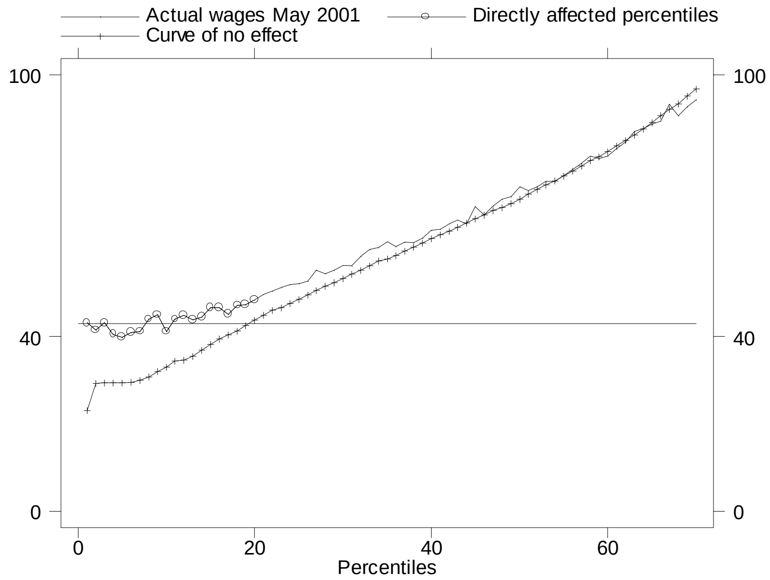

Figure I: Average wages in May 2001 in the 1st-70th percentiles of the May 2000 wage distribution

The wage data relates to 52,057 full-time employees observed in the 2000 wave of the WS and identified in the 2001 wave on the basis of the employer's ID, the plant's location, and the worker's age, gender, education and four-digit occupational code. The percentiles directly affected by the minimum wage hike are marked with circles. The horizontal line denotes the May 2001 value of the new minimum wage (thousand Ft). The 'curve of no effect' assumes that wages grew by the product of GDP growth and inflation all along the distribution.

The figure compares actual wages in May 2001 to the May 2000 wages multiplied by the product of price inflation and real GDP growth - a benchmark predicting the rate of nominal wage growth almost perfectly in the upper tiers of the distribution. (See the 'curve of no effect' marked with plus signs). Wages apparently grew much faster than that in the 1st-21st percentiles comprising workers directly affected by the minimum wage hike while a minor, gradually fading spillover effect was at work in the 15th-40th percentiles.

[4] The data in this and the next paragraph relates to firms employing 5 or more workers and the public sector. Author's calculation was using the Wage Survey.

These patterns justify the use of a simple, slightly downwards biased, indicator of 'shock to the average wage'. Under full compliance and negligible spillover a firm or occupation's exposure to the minimum wage hike can be approximated as:

$$(1) \quad \omega = [w^M F + w_H(1-F)] / [(w_F F + w_H(1-F)]$$

with F denoting the fraction of workers paid below the new minimum wage, w_F being their average wage at the moment of the hike, w_H standing for the average wage of other workers and w^M denoting the new minimum wage.[5] The formula measures the average wage gap to be filled on the day of the hike under the assumption that all sub-minimum wages rise to the level of the new floor and there is no further instantaneous wage and employment adjustment. As such, ω is a hypothetical benchmark that does *not* measure the actual response of average earnings but serves as a useful tool for the study of actual evolutions.

The hike was estimated to cause an immediate shock of 2.3 per cent to the economy-wide average wage under the assumptions implicit in ω. Calculating exposure for the interactions of five age groups, three educational levels and four quartiles of the country's 150 micro-regions (by unemployment) we got that group-level exposure varied between 0.3 and 16.7 per cent while F varied between 5 per cent and 60 per cent.

Whether the indications of the *payroll data* quoted so far should or should not be taken at face value requires further inspection (*Table I*). There are several ways of non-compliance remaining unobserved in the official reports. *First*, employers may simply withdraw a part of the reported wage. This kind of misuse may not be wide-spread: self-reported survey data indicated 1.4 per cent share of sub-minimum monthly earnings in April 2001, a ratio close to the payroll-based estimate of 1.9 per cent (first and second rows). *Second*, firms may employ their workers full-time but register them as part-time to be able to pay sub-minimum monthly wages. Indeed, a slightly higher fraction (3.6 per cent) of the employees who *actually* worked 36 or more hours a week in April-June 2001 reported sub-minimum earnings in the Labor Force Survey (LFS).[6] The estimate of earnings below Ft 40,000 or its part-time equivalent was 3.3 per cent according to the same source. *Third*, firms may fraudulently lay off their workers and contract with them as subcontractors. This sort of manipulation also seems

[5] Since our wage observations related to May we spoke of sub-minimum wages if a worker's wage was lower than Ft 38,685, the new minimum wage discounted with wage inflation between May and January.

infrequent. According to a survey of unemployed workers finding jobs in April 2001 (EJS) only 1.5 per cent concluded a business contract with the employer as opposed to 64.7 per cent receiving a fixed salary and 33.8 per cent paid an hourly wage.

Table I: Compliance with the law – Selected indicators

		Source, date, unit of observation
Per cent paid below the new minimum wage		
Employees registered as full-time	1.9	WS, May 2001, payroll data
Employees registered as full-time	1.4	EJS, April 2001, self-reported
Employees actually working full time	3.6	LFS, April-June 2001, self-reported
All employees[1]	3.3	LFS, April-June 2001, self-reported
Per cent paid as subcontractor	1.5	EJS, April 2001, self-reported
Elasticities with respect to ω		
∂(base wage)/ ∂ ω	0.96	WS and LFS, May 2001/May 2000[2]
∂(earnings)/ ∂ ω	1.00	WS and LFS, May 2001/ May 2000[2,3]
∂(earnings+taxes)/∂ω	1.00	FR, 2001/2000[4]
∂(all payments to persons +taxes)/∂ω	0.95	FR, 2001/2000[4,5]

1) Paid below Ft 40,000 if full time or Ft (h/40)·40,000 if part-time with h denoting usual weekly hours in the respondent's main job.
2) OLS estimates from a model where the log changes of average base wages were regressed on ln(ω) and a dummy for higher education background using data on 60 groups formed by interacting age, education, and region (see the text and Kertesi and Köllő [2003a] for details).
3) Earnings include overtime pay, shift pay, and bonuses
4) 2sls estimates from a two-equation system composed of a wage equation (right-hand side variables were log change in productivity, fraction unionized, mean regional unemployment, and sector dummies) and an employment equation (log change of output, ln(ω), the share of small firms in the industry, and sector dummies). Wages, employment and hence productivity were assumed to be endogenous. The system was estimated for 432 industries. For details see Kertesi and Köllő [2003a].
5) Other payments include per diem, honoraria and casual pecuniary benefits payable to individuals who are *not* necessarily accounted as employees.

Fourth and most importantly, firms can increase the base wage and reduce some side payments exempt from the regulations. The pecuniary offsets, however, unveil in comparisons of base wages with broader concepts of worker compensation. Most side payments, particularly shift pay and overtime pay, are set as percentages of the base wage therefore regular monthly earnings are expected to rise at approximately the same rate as do base wages if firms comply with the regulations. As shown in the bottom panel of *Table I* the elasticities of earnings and labor costs with respect to ω (using grouped and industry-level payroll data) fell close to unity, reinforcing that the first minimum wage hike was effective.

[6] The bias from not distinguishing between base wages and earnings in this case is predictably minimal as these fall close to each other at the lower tail of the wage distribution. The average earnings and base wages of workers earning less than Ft 40,000 in May 2001 were Ft 35,025 and Ft 34,736 respectively. (WS).

Figure II: Employment growth before and after the minimum wage hikes

Left panel: seasonally adjusted monthly employment in the non-agricultural private sector, million, 1998-2002. Right panel: seasonally adjusted quarterly employment in the non-agricultural private sector and GDP, normalized to their 1997. Q4. levels. The vertical lines separate the years. Sources: LFS for employment, seasonally adjusted by the National Bank, and National Accounts 2003 for the GDP.

The readily available descriptive statistics furthermore suggest that employment was adversely affected by the wage shock. *Figure II* indicates a sudden break in the growth of aggregate employment as soon as January 2001. The path of employment growth in and after 1998 (the first year when the number of jobs was rising since the mid 1980s) could be precisely approximated with a quadratic form indicated by an unmarked curve on the left panel. Had the economy remained on this path, as depicted by the extrapolated part of the curve, aggregate employment should have grown further by 2.8 per cent in January-December 2001 as opposed to an actual decrease of 0.2 per cent. The picture does not change if we consider the path of employment relative to GDP. In and after 1998 the economy followed a path at which one per cent growth of GDP was associated with half per cent growth of private non-agricultural employment. As shown by the right panel of *Figure II*, even with the slow-down of economic growth employment should have risen by about 1.7 per cent in 2001 and 1.8 per cent in 2002 in case of no break in the path of growth.

Grouped data relating to the interactions of 5 age categories, 3 educational levels, and 4 quartiles of regions (*Table II*) furthermore suggests that wage levels and employment records were negatively correlated in 2001. This pattern was at odds with previous experience: the group level ω-s (as of January 2001) and employment growth were unrelated in 1998-1999 while in 1999-2000 the low-wage groups experienced a rise in their relative employment probabilities.

Table II : Employment and exposure to the 2001 minimum wage increase – Regressions using grouped data

	1998-1999		1999-2000		2000-2001	
	All	Unskilled	All	Unskilled	All	Unskilled
			OLS regressions			
ln(ω)	–0,0936	–0,3682	0,9986***	0,9987***	–0,5431**	–0,9566***
Age > 55	0,1601***	0,1572***	0,2574***	0,3294***	0,0596**	0,0626*
Constant	0,0086	0,0319	–0,0291	–0,0396	0,0096	0,0396
Nobs	60	40	60	40	60	40
			Robust regressions			
ln(ω)	0,2962	0,7803*	0,3809***	0,5619***	–0,5880***	-1,2490***
Age > 55	0,1381***	0,2693***	0,2019***	0,2437***	0,0253	0,0036
Constant	–0,0024	–0,0325	–0,0015	–0,0164	0,0182	0,0674
Nobs	60	40	60	40	59	39

Dependent variable: log change of the employment/population ratio. Employment is defined on ILO grounds and relates to the working age population less old-age pensioners and students in 1999-2001, and the working age population in 1998-99. (Due to change in the registration of students the definition used later was not applicable in 1998-99). For the definition of groups see the text. The dummies for the oldest age group control for the effect of increases in the mandatory retirement age. Robust regression is estimated to mitigate the effect of a few heavy outliers. *Data on employment*: LFS 1998-2001 fourth quarters.. *Data on exposure*: WS 2000. The groups are weighted with base period size. The null of all coefficients being zero is rejected in each equation. The parameters are significant at the *) 0.1 **) 0.05 ***) 0.01 level.

The descriptive statistics obviously do not identify the effect of the rising minimum wage – the observed changes may have been driven by unobserved wage-specific demand or supply side shocks. The forthcoming sections try to disentangle the impact of the minimum wage using disaggregated data.

III. EMPLOYMENT IN SMALL FIRMS 2000-2001

In this section we analyze the effect of exposure to the minimum wage hike on changes of employment between 2000 and 2001 using annual firm-level data. The detailed analysis is restricted to small firms for two reasons. First, about 70 per cent of the Hungarian minimum wage workers are employed in firms with less than 50 employees. Second, as a fortunate coincidence, for at least a part of small enterprises the WS provides an exceptionally rich set of matched employer-employee data. As opposed to the general sampling rule (firms are expected to provide information on ten per cent random samples of their employees) companies with 5-20 workers are randomly sampled and expected to provide data on *all* employees, allowing a precise measuring of exposure. The section starts with the analysis of changes in small-firm employment between 2000 and 2001. This is followed by a study of linkages between wage levels (hence exposure) and employment in other years and other size categories.

When the minimum wage is adjusted the actual changes of wages are expected to exert strong influence on employment given a truly exogenous variation in wage growth. This effect can be captured by conditional labor demand equations similar to (2) with *L*, *y*, *p* and *w* standing for employment, value added, sales prices, and average labor costs, while the ***X***-s control for supply and demand shocks not captured by Δy .

$$(2) \quad \Delta \ln(L)_i = \alpha_0 + \alpha_1 \Delta \ln(y/p)_i + \alpha_2 \Delta \ln(w/p)_i + \boldsymbol{\alpha}_3 \mathbf{X}_i + v_i$$

As far as ΔL and Δw are endogenously determined the OLS estimate of (2) is inconsistent. Machin et al. [2003] estimate an equation analogous to (2) by instrumenting Δw with their 'shock to the average wage' variable, a close relative to our ω, and treating other firm-level variables as exogenous. This is one of the specifications tested later.

The impact of ω on Δw can also be explicitly modeled and taken into account in several ways. As long as the wage is exogenously determined a wage equation similar to (3) with ω (or *F*) on the right hand can be substituted to (2) to estimate $\partial L/\partial \omega = \alpha_2 \cdot \beta_1$, a parameter capturing the combined effect of compliance and the wage elasticity of demand for labor. A simultaneous equations system composed of (2) and (3) has the advantage of coping with simultaneity (by treating output, employment and wages as potentially endogenous) and separating the effects of β1 and α2.

$$(3) \quad \Delta \ln(w/p)_i = \beta_0 + \beta_1 \ln(\omega)_i + \boldsymbol{\beta}_3 \mathbf{Z}_i + u_i$$

For a brief discussion of the difficulties arising when it comes to empirical work we re-write the equations as a system (4-5) with *P* standing for *industrial* sales prices and *G* denoting group affiliation to allow structural breaks in the effect of ω on wages. The parameters of this model are potentially subject to endogeneity bias, on the one hand, and errors in the measurement of prices and exposure, on the other.

While the small firms under examination are most probably price-takers within their industries, *industrial* prices themselves can be affected by industry-level exposure. We tested this by regressing *P* on *F* and ω measured on the four-digit industry level (as well as on the level of 32 groups with distinct values of *P*). All specifications and estimation methods suggested that price movements were unrelated to the level of exposure. The endogeneity of output, wages, and employment were analyzed using Durbin-Wu-Hausman tests that rejected the exogeneity of labor costs, but not of output, in the particular empirical specification chosen for (4-5). [7]

(4) $$\Delta \ln(w/P)_i = \beta_0 + \sum_j \beta_{1j} [\ln(\omega)_i \cdot G_{ij}] + \boldsymbol{\beta}_2 \mathbf{Z} + u_i$$

(5) $$\Delta \ln(L)_i = \alpha_0 + \alpha_1 \Delta \ln(y/P)_i + \alpha_2 \Delta \ln(w/P)_i + \boldsymbol{\alpha}_3 \mathbf{X}_i + v_i$$

The estimates may also be affected by at least two types of measurement error. First, a bias of unknown direction may stem from unobserved within-industry price shocks correlated with the level of wages. Let p_i and Pi stand for firm-level and industry-level prices so that $\Delta ln(p)_i = \Delta ln(P)$i+ξi , E(ξ)=0. Since sales and wages are discounted with *P*i rather than *p*i the residuals of (5) become εi=νi+(α1+α2)ξi as opposed to *v*i in equation (2). For ω to be a valid instrument E(εω)=0 is required, which may not be the case if the within-industry variations in price movements are correlated with the level of wages and hence ω.[8] Though the Hungarian economy was free of major shocks until after 9/11/2001 this sort of bias may be present in the estimates.

Errors in measuring *exposure* have potentially more severe implications. The bias stems from the fact that some workers registered as minimum wage workers are paid additional remuneration in cash. The costs of employing such workers increased by the difference between taxes levied on the old and new minimum wages – far less than 57 per cent. To the extent these practices prevail ω overestimates the magnitude of the minimum wage shock, and the predictive power of *F* declines. The degree of parameter bias depends on the correlation between *F* and the share of 'genuine' minimum wage workers within *F*. We tried to ascertain this sort of correlation by estimating equation (3) with *F* on the right hand side, with both OLS and IV, using variables on the firm's skill composition as instruments. The coefficients were robust to changes in the method of estimation suggesting that 'under the counter payments' do not strongly affect the parameter estimates.

Sample, data, and empirical specification. The sample was drawn from the population of enterprises interviewed in the 2000 wave of the WS. In each cross-section wave small firms

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[7] The test statistics are presented together with the estimation results. With output treated as exogenous the system passes both the overidentification and the exclusion restrictions tests allowing the estimation with 3sls.

[8] This is less of a problem in the Machin et al. [2003] paper since they analyze a homogeneous sample of residential care homes at the time the minimum wage was reintroduced in the UK.

are randomly selected within strata formed by four-digit industries. Given the target population of small firms and the sampling quota the expectation was that about 350 small firms could be followed in a short panel out of the 2,874 companies interviewed in 2000. In fact, the number of small enterprises observed in 2000 and 2001 amounted to 2,008. This regrettably calls into question the alleged independence of the cross-section samples but fortunately provides us with a sizable longitudinal sample drawn from a populace of firms heavily exposed to the minimum wage shock. Out of the 2,008 firms 1,818 had all the variables required for the estimation.

Table III: Small firm panel 2000-2001 - Probits of sample selection

Sample	Dependent variable = 1	Number of employees	Fraction low-wage	Lossmaker in 2000	Pseudo R2	Nobs
Small firms observed in 2000	Also observed in 2001	.0012 (2.43)	-.1074 (4.96)	-.1239 (5.93)	.0209	2,874
Small firms observed in both 2000 and 2001	Has complete data	.0036 (2.51)	-.0099 (0.60)	-.0581 (3.17)	.0166	2,008

*) The table shows the marginal effects

The probits in Table III check how the estimation sample was selected from the base-period population of firms. The companies also observed in May 2001 were larger, generated profit in the base period; and had fewer workers paid below the new minimum wage. The dropouts were predictably hit harder so our models underestimate the extent and potentially adverse implications of the minimum wage shock. The estimation sample within the panel is also biased for larger firms and profit makers but does not systematically differ from the rest of the sample in terms of exposure.

The data on annual average employment, annual average labor costs (all payments to individuals plus social security contributions), and output (sales revenues net of material costs and depreciation) were taken from the firms' annual financial reports (FR). The descriptive statistics of the estimation sample are presented in the Appendix. The median firm had 13 employees of which 5 was paid below the new minimum wage, and was hit by an average wage shock of 11.2 per cent.

Table IV: Small firms – Performance in 2000-2001

Characteristics in 2000 per cent			Mean log change 2000-2001 weighted with base period employment				Number of firms
Brackets by Minimum wage shock	Fraction low-wage	Mean ω	Average wage	Labor cost[1]	Output[1]	Employ-ment	
0	0	0	.121	.063	-.009	-.004	468
0-10	27.4	3.2	.154	.088	-.026	-.025	632
10-25	74.1	16.6	.274	.172	-.038	-.081	319
> 25	95.9	35.8	.398	.309	-.017	-.105	399
All firms	43.5	11.3	.216	.141	-.022	-.043	1,818
Anova[2]			117.2	87.8	0.3^n	6.8	

1) Discounted with industrial producer prices (32 distinct values). 2) F-test for the equality of means. Equality is rejected at the .001 level except for output (rejected at the .833 level)

In the estimated specification of system (4-5) the uniformity of the wage effect of ω across regions was tested under the assumption that the level of compliance was higher in depressed regions, where failures to pay the new minimum wage would have menaced with the quitting of core workers. Four groups of micro-regions were distinguished by unemployment (Employment was assumed to respond to output and wages uniformly across regions as suggested in Kőrösi [2000]). Equation (5) included the base period capital-labor ratio under the assumption that capital intensive firms were less likely to react with dismissals on the short run. Equation (4) included base period profits to allow for the effect of profit sharing. In the employment equations 18 region dummies controlled for supply shifts and 10 industry dummies allowed for demand shocks unobserved in Δy, and changes in technology.

Results. Table IV gives a descriptive overview of changes between 2000 and 2001, broken down by the magnitude of the minimum wage shock. Real labor costs grew and employment fell sharply as a function of exposure.

Table V: Estimates of equations (2) – (6)

Variables:	3SLS[1] (1)	IV (ω) (2)	IV (F) (3)	OLS (4)	OLS (5)	OLS (6)	OLS (7)	OLS (8)
Output (Δln)	.2522***	.2709***	.2709***	.2494***	.2501***	.2468***	.2486***	.2459***
Labor cost (Δln)	-.4089***	-.4010***	-.4031***	-.0061	-	.0617*	-	.0464
MW gap (ln)	-	-	-	-	-.2913***	-.3356***	-	-
Fraction affected	-	-	-	-	-	-	-.0958***	-.1069***
Controls[2]	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Constant		.0757*	.0762*	-.0017	.0179	.0091	.0249	.0004
F (χ^2 for 3SLS)	140.1***	6.42***	6.30***	6.21***	6.74***	6.83***	6.60***	6.65***

Significant at the *) 0.1 **) 0.05 ***) 0.01 level. **1)** The coefficients of ln(ω) in the wage equation are .6554, .7071, .7629 and .7703 in the first-fourth quartiles of micro-regions, respectively. The coefficient of base period profits is .0003. All coefficients are significant at the .001 level. Specification tests. Durbin-Wu-Hausman for endogeneity: P>|t|=0.001 for labor costs, 0.272 for output. Sargant's overidentification restriction test: P(χ^2)= 0.051. F-test for the joint significance exogenous regressors excluded from the employment equation: 0.002. **2)** The controls include the base period capital-labor ratio, 10 industry dummies, and 18 region dummies

The estimates of the OLS, IV and 3SLS models with ω and F used as alternative measures of exposure are presented in *Table V*. The wage setting equation of the 3SLS (summarized in the bottom row of the table) suggested that the elasticities of labor costs with respect to the minimum wage shock ranged between 0.66 and 0.77 with high-unemployment regions having higher values. Generally, we found lower levels of compliance than in Section II where grouped or industry-level data were used.

The elasticities of *employment* with respect to output varied in a narrow range of 0.25-0.27. This finding is consistent with an estimate of 0.3 in Kőrösi's [2002] differenced Cobb-Douglas model using firm-level data for 1996-99. The wage elasticity of labor demand appears to be about –0.4, also consistent with Kőrösi's estimates averaging to –0.3. From the 3SLS estimates of $\partial w/\partial\omega$ and $\partial L/\partial w$ we can predict $\partial L/\partial\omega$ to vary between –0.28 and –0.31 depending on region while the OLS estimate for all firms is -0.29 in column 5.

In the OLS model ignoring the information on exposure (column 4) the wage elasticity estimate is insignificant reflecting strong attenuation bias. Adding ω or *F* to the equations (columns 6 and 8) results in insignificant positive coefficients for Δw, and highly significant negative estimates for ω and *F*, reinforcing that employment was affected by variations in exposure rather than variations in Δw at given levels of exposure.

Contrast with previous experience. If low wages observed in one year are generally conducive to employment cuts in the next year this linkage is captured as a `minimum wage effect' in our models. Indeed, low wages may result from poor firm performance indicative of forthcoming employment cuts, or signal lags in the process of wage adjustment so that the periods of low wage levels are followed by periods of fast wage growth and employment cuts. The results in *Table VI* call into question if such a general rule applies to the Hungarian small firm sector. Changes of employment were unrelated to the level of wages and the share of low-wage workers in 1999-2000 unlike in the period of the minimum wage hike.[9]

Table VI: Base period average wages and employment growth - Univariate regressions using data on small firms

Dependent variable: log change of employment	1999-2000	2000-2001
Base period log average wage	-0.014	0.056***
Fraction low-wage in 2000 (w<Ft 38,685 → 1.-21. Percentiles)		-0.121***
Fraction low-wage in 1999 (1.-21. percentiles → w<Ft 34,953)	0.004	

***) significant at the 0.01 level, unmarked coefficients are not significant at the 0.1 level. Data source: FR and WS1999, 2000, 2001. Number of firms 1.046 in 1999-2000 and 1,818 in 2000-2001.

Larger firms. The information available for a similar analysis of larger firms is less reliable because these firms report individual data on ten per cent random samples of their employees. The obsrevations on F or ω from these small samples are noisy but not systematically biased. Repeating the estimation of the IV model with ω used as the instrument, and the OLS model with ω on the right hand, yields the elasticities reported in *Table VII*.

Table VII: Estimates for all firms observed in 2000-2001 (WS)

Firm size	Mean exposure		Elasticities of employment with respect to			Number of firms	Employment	
	F	ln(ω)	Output [1]	Labor cost [1]	ω [2]		Panel [3]	Total [4]
5-20	43.5	.113	.2708	-.3932	-.2909	1,818	229,523	342,804
21-50	30.8	.080	. 2289	-.4186	-.2076	2,555	136,052	180,076
51-300	18.7	.044	.3489	-.4307	-.2114	2,846	375,614	449,065
301-	7.0	.013	.7517	-.0421[n]	-.0561 [n]	572	676,362	748,899

n: Not significant at the 0.1 level. Unmarked coefficients are significant at the 0.01 level. 1) Specification 2 of *Table V* 2) Specification 5 of *Table V*. Weighted with base-period employment. 3) Aggregate employment in the firm panels. 4) Target population of the WS of May 2000.

[9] Data for firms employing 5-10 workers are only available since 1999. The short panels built for firms with 11-20 workers in 1997-98 and before are too small for a similar kind of analysis (contain only about 100 firms).

Output elasticities fall and wage elasticities increase in absolute value as we move from small to large enterprises. As a combined effect of lower wage elasticities and lower levels of exposure the implied employment losses become smaller with medium sized firms and virtually zero with large firms: the estimates are -3.5, -1.6, -0.8 and -0.1 per cent in the four size categories, respectively. Since large firms account for a considerable part of private sector employment we also get a relatively low estimate of -1.1 per cent for aggregate employment loss in the WS target population that excludes the public sector, sole-proprietors and firms with 1-4 employees.

The lower bound estimate for the whole economy assuming no minimum wage effect in the excluded sectors is -0.5 per cent. For an upper-bound estimate one should first consider that the value of ω was only 1.5 per cent in the public sector therefore the implied employment loss could be easily averted by marginally higher budget expenditures. The effect on sole proprietors must have been negligible, too, as they could easily evade the regulations. Firms employing 1-4 persons are similar to those with 5-20 workers in terms of wage distribution, and have a similar share in aggregate employment. Assuming similar exposure and response, and adding the implied loss of jobs to what we have from *Table VII*, we get an upper-bound estimate slightly below one per cent for the whole economy.

IV. IMPACT ON LOW-WAGE WORKERS

Losing one out of 100 (or 200) jobs may seem to be a negligible price paid for a 57 per cent rise in the minimum wage that helped to increase the earnings of one in five workers. This, however, is the median voter's view of the trade-off – an aspect becoming less and less relevant as we move toward 'lower' segments of the labor market. This section provides information on how low-skilled and/or low-wage workers were affected. Data availability does not allow a comprehensive overview but we do have pieces of meaningful information from areas severely exposed to the risks of *in vivo* experimentation with the minimum wage.

IV.a. Employment by skills in small firms 2000-2001

First, we briefly return to our panel of small firms to benefit from the repeated cross-section information on individual employees. Since the workers can not be identified across waves we can not observe the wage-specific changes in employment. However, the percentage changes in the share of low-educated and blue-collar employees (*Table VIII*) clearly show that the burden of adjustment fell on the low-skilled workers of low-wage firms. While the share of low-skilled labor increased, remained constant, or fell marginally in high-wage firms, it markedly declined in firms strongly exposed to the minimum wage hike.

Table VIII: Change in the percentage share of low-skilled labor in small firms 2000-2001

Type of labor	Weighting[3]	Fraction low-wage in the firm (per cent)				Total
		0	0-10	10-25	25-100	
Low-educated[1]	Yes	1.2	0.0	-4.2	-2.9	-1.4
	No	-0.2	.1.5	-2.2	-2.0	-1.4
Blue-collar[2]	Yes	3.9	0.3	-0.1	-1.9	0.3
	No	1.2	-0.6	-2.0	-3.0	-0.9

1) Lower than secondary education (incomplete primary, primary, uncertified vocational). 2) According to the worker's four-digit occupational code 3) With base period employment. The data relate to the 1,818 small firms analyzed in Tables IV-VII.

IV.b. The jobloss risks of low-wage wage workers, March-December 2001

A minimum wage hike decided in a government office randomly divides the low-wage population into two parts. Workers whose pre-hike wages were just above the new minimum are likely to have similar human capital endowments and occupational characteristics to those who earned just below the line but their employers have no straightforward motivation to fire them as they are kept to be paid at their marginal products. These workers can also be indirectly affected by wage spillovers or because the firm's demand falls for the whole category of labor they belong to. Still there is likely to be a difference in the jobloss probabilities of those directly affected and those who are not, or only indirectly, influenced. Following this line of reasoning we study how wages affected the jobloss hazards of full-time employees interviewed in the LFS Supplementary Survey of 2001 2nd quarter. [10]

We distinguish a *treatment group* (workers who were paid the new minimum wage) from a *control group* (those who earned slightly more than that) and estimate the two group's jobloss probabilities in March-December 2001 using a discrete time duration model.[11] Our approach is similar to that of Currie and Fallick [1996] and Abowd et al. [1997] both comparing workers paid the minimum wage with those earning just above the limit.

Sample restrictions. Workers in marginal jobs change employer frequently so they tend to have high jobloss probabilities and low wages at any point in time. In order to minimize the

[10] This was the only wave since 1993 when respondents were asked about wages in the LFS.
[11] As shown in Jenkins [1995], by choosing the quarterly employment spells of individuals as the units of observation the exit hazard from a stock sample can be estimated with logit augmented with a baseline hazard function.

influence of this correlation we restrict the attention to workers who spent at least two years in their jobs prior to the survey date. (The treated and the controls spent 6.7 and 7.3 years in their jobs on average.) Workers were followed by the end of 2001. The reason of not following them for 5 quarters, the longest possible period allowed by the LFS design, is that the second minimum wage shock exposed the control group to the same type of risk that hit the treatment group in 2001. The analysis is restricted to full-time employees. After these restrictions the estimation sample contains 22,315 quarterly employment spells.

Wage brackets. The wage data relate to gross monthly earnings as reported by the respondents, or estimated from the net figure by the CSO. We distinguished workers paid 90-110 per cent of the minimum wage (*treatment)* from those earning 110-125 per cent (*control*), and three other categories earning higher wages.[12] The brackets were chosen to maximize the distance between the treatment and control groups in terms of exposure to the minimum wage increase according to data from the WS Individual Panel of 2000-2001 introduced earlier. The estimate is that 83.6 per cent of the treatment group was likely affected but only 54.4 per cent of the controls were unaffected. Since the vast majority of the misclassified workers are found in the control group the model underestimates the treatment effect [13]

Results. There was a large and statistically significant difference between members of the treatment and control groups in their probability of becoming *unemployed* in the 2nd -4th quarters of 2001 as shown by the coefficients of 1.05 versus 0.15 significantly different from each other at the 0.04 level (*Table IX*). While the exit to *non-participation* hazards were equal in the two groups minimum wage workers were more likely to lose their jobs *and* try to get back to work through active job search.

[12] Workers earning less than Ft 36,000 were excluded from the estimation sample because this category apparently includes many workers planning to retire. Furthermore, we observed high wage mobility between this and other brackets suggesting that sub-minimum wages are often explained by temporary reasons.

[13]. It might also be mentioned at this point that the second minimum wage hike that became a credible promise/threat by the autumn of 2001 also biases the observed treatment effect downwards.

Table IX: Exit from employment 2001 2nd-4th quarters - Discrete time duration model, multinomial logit form

	Left employment for			
	Unemployment		Non-participation	
Male	-.0948	-0.31	-.5615	-3.10
Age	.5116	3.39	-.3338	-6.75
Age squared	-.0063	-3.38	.0041	7.01
Unskilled blue collar.	-.1559	-0.32	-.4750	-1.20
Semi-skilled blue collar	.1277	0.33	.0850	0.34
Skilled blue collar	.2456	0.64	-.0048	-0.02
Unemployment (log)	-.01664	-0.08	.3708	2.54
Public sector	-.9144	-1.65	-.0598	-0.22
Union member	-.7295	-1.82	.1420	0.63
Tenured job	-.3427	-0.62	-.6559	-2.08
Wage Ft 36,000-44,000 (treatment)	1.0596	3.00	.1078	0.44
Wage Ft 44,000-50,000 (control)	.1494	0.31	.0600	0.19
Wage Ft 75,000-100,000	-.5536	-1.14	-.4572	-1.63
Wage Ft >100,000	-.0494	-0.10	-.3114	-0.97
2001 4th quarter	.3108	1.09	.3152	1.79
Exp (-tenure in years)	4.4246	2.61	-.2657	-0.09
Constant	-15.5637	-5.06	2.8677	2.50
Coefficients from an alternative specification:				
Wage Ft 36,000-44,000 (treatment) * U	3.9671	2.13	3.6431	2.37
Wage Ft 44,000-50,000 (control) * U	-1.3663	-0.38	2.3481	1.27
Wage Ft 50,000-75,000 * U	-3.8709	-0.93	3.9035	2.68
Wage Ft 75,000-100,000 * U	-10.578	-1.56	-.76228	-0.29
Wage > Ft 100,000 * U	-8.7554	-1.55	3.2095	1.20

Logit coefficients and Z values. Reference categories are white collars, wage Ft 75,000-100,000, tenure>18 months. Test statistics of the base specification. Number of observations: 22,315. -log likelihood: 1302.12. Pseudo R^2: .0525. F-test for the equality of the coefficients of the treatment and control groups: 4.13 (.0421) in the unemployment equation and 0.02 (.8906) in the non-participation equation. Standard errors adjusted for clustering by individuals. Data sources: LFS 2001 2nd quarter Supplementary Survey, LFS 2001 3rd and 4th quarters.

The estimated quarterly outflow to unemployment rates of 25 year old male workers with 5 years of tenure were 0.243 and 0.119 per cent in the treatment and the control groups, suggesting rather long prospective tenures. The fraction *not* becoming unemployed until retirement is estimated to be 67.5 and 82.6 per cent in the control and treatment groups, assuming a retirement age of 65 and constant hazard.[14] The exit to unemployment hazards increased with regional unemployment *within* the minimum wage group while at higher wages the regional differences were negligible. (The equality of the coefficients can be rejected only at the 0.09 level. The parameters for exit to non-participation are statistically equal.) This is consistent with the finding of a stronger minimum wage effect at small firms located in high-unemployment regions.

[14] The estimation results are qualitatively similar assuming constant or piecewise baseline hazard by including a years of tenure variable in the first case, and dummies for 3-5 years of tenure in the second. The small number of exits did not allow flexible baseline hazard with dummies for longer tenures.

IV.c. Outflows from unemployment of low-wage workers, 1998-2002

The orthodox wisdom predicts a fall in the job finding probabilities of unemployed workers who were paid below the new minimum wage prior to losing their jobs. Whether the outcome was predominantly shaped by a classic demand-side effect, or by more complex mechanisms offsetting the adverse implications of the minimum wage shock, is tested using a panel of 172 labor offices observed between January 1998 and June 2002.

For each office and month we know the number of low-wage and high-wage workers among unemployment insurance benefit (UI) recipients at the beginning of the month and their exit to job rates (hLW and hHW) during the month. The same information is available for low-skilled and high-skilled workers (hLS and hHS). The return to comparing low-wage and high-wage workers is clearly minimal as these groups largely differ in terms of skill levels and exposure to economic shocks. In order to get closer to a sensible comparison we study how the exit rates of low-wage workers related to the exit rates of low-skilled workers before and after the minimum wage hike. The approach is closest to that of Deere et al. [1996] analyzing teenage employment after increasing the US federal minimum wage. We estimate equation (6):

$$(6) \quad \ln(h^{LW})_{it} = \beta_1 \ln(h^{LS})_{it} + \beta_2 \ln(U)_{it} + \boldsymbol{\beta_3}\mathbf{MD} + \boldsymbol{\beta_4}\mathbf{YRD} + c_i + v_{it} \ ,$$

where h_{it} is the exit rate at office *i* month *t*, *LW* and *LS* refer to low-wage and low-skilled workers, and *MD* and *YRD* are month and year dummies. The long-run averages of the office-level *h*LW/*h*HW ratios may differ depending on the typical duration of unemployment of the low-wage and unskilled groups – the resulting time-invarying fixed effects are captured by the ci-s.[15] The expectations are β_1=1 and $\beta_2 \leq 0$ (as it is more difficult for low-wage workers to find jobs when the market is depressed). Prior to the minimum wage hike the year effects are expected to fall close to zero but a significant break is anticipated in 2001.

The equation has to be instrumented for obvious endogeneity on the one hand, and possible correlation between the residual and *h*LS on the other. Some sort of regional shocks may exert strong impact on *h*LW relative to *h*LS. When whole plants are closed or opened employers

[15] The mean benefit divides the population of UI recipients to fractions of varying size depending on the regions' wage level. The difference in the skill endowments of the median recipient and the median low-wage recipient

screen their workers/applicants more carefully and while doing so they interpret low-wages as a signal of low productivity – this establishes a link between *h*LS and *v*it. The sign of the correlation is *a priori* unclear since *h*LW is expected to rise *less* when *h*LS is rising, and fall *more* when *h*LS is falling. We instrument *h*LS with its *t-1* period value.[16]

Measuring low skills and low wages. The labor offices record the recipients' earnings in the four calendar quarters preceding their current unemployment spells. Since the benefits are earnings-related they also provide an indirect measure and we use them as a proxy of the wage. (Though pre-unemployment earnings are known they relate to different time periods - computing the present value of past earnings case by case would have enormously increased the costs of data collection.) Data from the UI register of March 2001 show that the benefit is indeed a good proxy of the wage: 98.7 per cent of the workers receiving lower than average benefits earned less than the median wage prior to unemployment, and 87.2 per cent of the high-benefit recipients had higher than median wages. Altogether, 92.3 per cent of the recipients could be correctly classified as 'low-wage' or 'high-wage' on the basis of their benefits. Skilled workers are those with completed secondary and higher education. The available data suggest that 81.4 per cent of the low-wage workers were low-skilled but only 48.8 per cent of the low-skilled were low-wage therefore hLW/hLS can be considered a crude approximation of the wage-level specific job finding rate (hLW|LS) within the unskilled group.

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tends to be smaller in low-wage regions, which provides an explanation for the regional fixed effects. Regional differences in the share of seasonal low-wage industries add a further component to c_i.

[16] Further complications might arise from the fact that the composition this month's inflows have an impact on the composition of next month's stock. We neglect this feedback because job finds account for less than 1/3 of the total outflows from the UI stock and the latter is also affected by the inflows. It is also worth noting that there is no straightforward link between the flows of the UI system and unemployment. In 2000 less than 20% of the ILO-unemployed received UI.

Table X: The exit to job rate of low-wage UI recipients 1998- 2002 – Panel estimates

	Fixed effects instrumental				Fixed effects	
	Missing values replaced[1]		Cases with missing values dropped		Missing values replaced	
Base specification						
Log exit rate of the low-skilled	1.0242	17.13	0.9560	15.96	0.8120	105.51
Regional unemployment rate	-0.0191	0.64	-0.0224	0.82	-0.0444	2.70
1999	-0.0199	1.80	-0.0199	1.97	-0.0274	2.68
2000	-0.0062	0.48	0.0051	0.41	0.0267	2.59
2001	-0.0883	5.88	-0.0742	5.26	-0.0451	4.36
2002	-0.1173	6.56	-0.0960	5.83	-0.0712	5.43
Constant	-0.0150	0.007	-0.2346	1.08	-0.5988	21.79
Alternative specifications:						
(i)						
2001-2002 dummy	-0.0871	8.42	-0.0778	8.07	-0.0536	7.35
(ii)						
2001-2002 × 1st quartile	-0.0863	5.31	-0.0782	5.37	-0.0589	4.12
2001-2002 × 2nd quartile	-0.0548	3.15	-0.0563	3.60	-0.0441	3.09
2001-2002 × 3rd quartile	-0.0967	5.69	-0.0873	5.63	-0.0605	4.18
2001-2002 × 4th quartile	-0.0992	5.21	-0.0819	4.58	0.0521	3.59
Tests of the base specification						
Within R2	0.7190		0.7363		0.7409	
Overall R2	0.7773		0.7846		0.7818	
Number of observations	9116		8975		9116	
Wald χ^2 (F for the FE model)	738744		890437		1502.44	
F-test for β_1 being unity	0.16	0.6857	0.47	0.4909	591.96	0.0000

Panel estimates using monthly data from 172 labor offices, January 1998 – June 2002. Dependent variable: log exit rate of the low-wage recipients. In 2 per cent of the cases the exit rate of low-wage workers were zero – these cases were excluded or the zeros were replaced assuming the outflow of ½ person. The coefficients of the month dummies are omitted.

Results. The estimation results of equation (6) are shown in Table X . In the fixed effects model β1 falls short of unity, a clear indication of attenuation bias, while in the IV-s they do not significantly differ from the expectation of β1=1. When unemployment increases the relative exit probability of the low-wage recipients falls but this effect is not significant at conventional levels. The month effects (not displayed) hint at changes in the composition of the low-wage unemployed pool over the year.17

Most importantly, the results indicate a 7-9 percentage points fall in the job finding probability of the low-wage unemployed relative to the unskilled in 2001, and a further 2-3 percentage points decline in January-June 2002. Testing the pair-wise equality of the year effects suggested that the parameters for 1998-2000 were not significantly different from zero and each other; those for 2001 and 2002 were strongly different from zero and any of the previous year effects, while they differed from each other at the .95 but not the .99 level of

[17] During the fall and winter when unskilled job opportunities are scarce and many young, low-wage unemployed return from `unemployment holiday` h^{LW} rises relative to a falling h^{LS}.

significance.[18] Treating the pre- and post-hike periods as different regimes when estimating the same equation with a dummy for 2001-2002 provided a coefficient of -.087. Interacting this 'regime dummy' with dummies for the four quartiles of regional unemployment (treating all regions in 1998-2000 as the reference) yielded statistically equal parameters.[19]

V. SUMMARY

The Hungarian government's decision to radically increase the minimum wage implied a loss of employment opportunities. Similar to the experience of Indonesia – the country providing the closest analogue to Hungary's minimum wage experiment – the data indicated minor impact on medium-sized and large firms. Our estimate of the short-run damage to aggregate employment fell between 0.5 and 1 per cent, implying an elasticity of aggregate employment with respect to the minimum wage somewhere between –0.01 and –0.02 – virtually zero. This is, in fact, close to what one could expect under a 2.3 per cent increase of the average wage (implied by the minimum wage hike) and demand elasticity estimates from other papers.

One might argue that even a loss of this magnitude can be painful for an economy destroying jobs for more than a decade, and expanding employment by less than one per cent annually in a short period preceding the minimum wage hike. In evaluating the impact, however, it seems more important that the severe implications are concentrated in narrow strata of the labor market: low-wage firms and low-wage workers were strongly hit already in the short run. The small firm sector lost at least 3.5 per cent of its jobs in less than a year, and the job retention and job finding probabilities of low-wage workers markedly deteriorated.

The finding that depressed regions were equally or more severely hit by the hike underlines the relevance of the classic framework in predicting minimum wage effects. The workers of high-unemployment regions have higher probability of receiving unemployment benefits; the benefits replace a larger fraction of their lost earnings; they have better than average access to informal second jobs, and are more severely constrained by fixed costs like travel-to-work expenses. The positive supply-side effects predicted in several theoretical models of the minimum wage, and envisaged by the Hungarian government, are more likely to develop under such conditions. The evidence suggests that, even in these regions, the positive

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[18] These results are available in Kertesi and Köllő (2003a,b,c)

[19] While the fixed effects capture the long-term differences in h^{LW} relative to h^{LS} they do not control for regional variations in the *changes* of the two exit rates, in response to a wage shock. In low-wage region more unskilled workers are paid low-wages therefore h^{LW}/h^{LS} changes little when h^{LW} falls. In high-wage regions a wage-related shock affects h^{LW} stronger than h^{LS} so h^{LW}/h^{LS} falls substantially. This leads to an underestimation of the effect hitting the low-wage regions.

responses were more than offset by the elementary cost effect of move to a higher minimum wage.

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Data appendix

FR – Financial Reports. The Ministry of Finance collects the reports of enterprises. The sample used in this paper is restricted to firms observed in the WS. The reports include an account of assets and liabilities and annual intakes and costs. The firms can be identified across waves. The descriptive statistics of the small firm panel are presented in Table A.I.

Table A.I.: Descriptive statistics of the small firm panel (N=1818)

Variable	Mean	Median	Standard deviation
Employment 2000	12.7	13	4.44
Employment 2001	13.6	12	14.30
Value added 2000 (mFt)	227.5	91.5	712.3
Value added 2001 (mFt)	251.3	98.0	891.6
PPI 2000-2001	1.066	1.063	0.025
Average wage 2000 (mFt)	0.824	0.583	0.901
Average wage 2001 (mFt)	0.978	0.700	0.992
Profit 2000 (mFt)	3.27	1	38.3
Assets/worker (mFt) 2000	4.816	1.333	29.1

WS – Wage Survey. The WS is an annual survey conducted by the National Labor Centre (NLC) each May since 1992. In the waves used in this paper the sampling procedure was the following (i) the firm census provided by the CSO serves as the sampling frame (ii) it is a legal obligation of each firm employing more than 20 workers to fill in a firm-level questionnaire and provide individual data on a 10 per cent random sample of the employees. (iii) budget institutions irrespective of size have to fill in the institution-level questionnaire and provide individual data on all employees (iii) Firms employing less than 20 workers according to the census are sampled in a procedure stratified by four-digit industries. The firms contacted are obliged to fill in the firm-level questionnaire and provide individual demographic and wage data on all employees. The observations are weighted to ensure representativity. About 180 thousand individuals employed in 20,000 firms and budget institutions were observed in 1999-2001.

LFS – Labor Force Survey. The LFS is a representative quarterly household survey conducted by the Central Statistical Office (CSO) since 1992. Data are collected about each member of the surveyed households and an ‘activity questionnaire’ is filled with those aged 15-74. The survey has a rotating panel structure with each quarter 1/6 of the sample dropped after spending 6 quarters in the survey, and replaced with a randomly chosen new cohort. The number of observations varied between 82 and 85 thousand in 1999-2001. Individuals can be identified across waves. The cases are weighted by the CSO to ensure representativity. All calculations in this paper used these weights.

Table A. II. : Jobloss - Descriptive statistics of the estimation sample (22,315 spells)

	Mean	St. dev.
Exit to unemployment	0.30	
Exit to non-participation	0.73	
Male	.5247	
Age	40.27	10.36
Unskilled blue collar	.0759	
Semi-skilled blue collar	.1689	
Skilled blue collar	.3502	
Regional unemployment	.0925	.0597
Public sector	.1741	
Union member	.2502	
Tenured job	.9617	
Wage Ft 36,000-44,000	.1522	
Wage Ft 44,000-50,000	.0932	
Wage Ft 75,000-100,000	.1919	
Wage Ft >100,000	.1718	
2001 4th quarter	.4344	
Tenure in job (years)	7.29	2.87

LFS Supplementary Survey April-June 2001. The LFS does not collect wage data. In this particular wave respondents working as employees or cooperative members (22,415 out of 30,485 workers employed by ILO-OECD standards) were asked to tell their last month's gross or net earnings. The gross value of net earnings was calculated by the CSO using tax tables. We used the gross figures as reported by the CSO and weighted the cases followed in a spell panel with their base period weights of April-June 2001.

NLC Office-level Exit to Jobs Panel 1998-2002. The data base was built in the NLC in September 2002 using data from Hungary's 175 labor offices. It contains aggregate stock and outflow to jobs data broken down by three levels of education (primary or lower; vocational; secondary and higher), and the level of the benefit (lower/equal or higher than the national mean). The stock figures relate to the first day of the month and the flows relate to the month. Three offices were involved in reorganization during the period of observation and were dropped from the sample analyzed in this paper. The unemployment rates attached to the offices are ILO-OECD counts divided by the population of working age, as estimated by the CSO, in the territory of the office. Job finds exclude entry to public works and other programs for the unemployed.

NLC EJS – National Labor Centre Exit to Jobs Survey, April 2001. The NLC interviewed all workers leaving the UI register because of job finding between March 22 and April 7,

2001. The workers were interviewed when they contacted the office to collect the documents necessary to enter employment. They were asked about their minimum and maximum expected gross monthly earnings in the first months after being hired. The file used in this paper contains the data of 105,957 recipients in the stock on 22 March 2001 and interviews with 9,131 workers finding a job. Of them, 8,811 workers provided wage data. The wage and benefit concepts used in the paper are (i) gross monthly earnings in the four calendar quarters prior to the last UI spell, adjusted for wage inflation between the time of jobloss and March 2001. (ii) The mean of the minimum and maximum expected earnings (iii) the monthly values of the pre-tax daily UI benefit assuming 30.5 days a month.

The Wage Effects of Schooling under Socialism and in Transition:

Evidence from Romania, 1950-2000*

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Abstract

We estimate the impact of schooling on monthly earnings from 1950 to 2000 in Romania. Nearly constant at about 3-4% during the socialist period, the coefficient on schooling in a conventional earnings regression rises steadily during the 1990s, reaching 8.5% by 2000. Our analysis finds little evidence for either the standard explanations of such an increase in the West (labor supply movements, product demand shifts, technical change) or the transition-specific accounts sometimes offered (wage liberalization, border opening, increased quality of education). But we find some support for institutional and organizational explanations, particularly the high productivity of education in restructuring and entrepreneurial activities in a disequilibrium environment.

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1. Introduction

An increasing number of studies have begun to document the rapid rise in relative earnings associated with education in post-communist Eastern Europe (see, e.g., the summary in Fleisher, Sabirianova Peter, and Wang, 2004). Little attention, however, has been paid to the schooling premium in Romania, the topic of this paper. The single available set of previous estimates, in a recent article by Skoufias (2003), pertains to only one cross-section of data in the early transition year of 1994. While these results may be compared with those from other countries, clearly they are limited in their ability to track the impact of transition, as they contain information neither on the pre-transition situation nor on developments as transition progressed through the 1990s. Indeed, many of the studies of earnings differentials in other transition economies are similarly limited to cross-sections or very short time series, and relatively few have analyzed databases with long series of information both before and after the tumultuous changeovers in political-economic system.[1]

In this paper, we use data from 1950 to 2000 to estimate the evolution of the wage impact of schooling for Romanian workers. Romania provides a particularly interesting setting in which to investigate these issues. To an even greater extent than in most other transition economies, Romania's economy during the socialist period up to 1990 reflected a thorough system of central planning and administrative controls, with none of the partial reforms adopted in Hungary, China, or the former Soviet Republics. Labor issues were strictly under the purview of the State Planning Committee, emigration was virtually prohibited, and migration was very strictly controlled, with 10 cities "closed" to new residents. Base wages were prescribed by the Wage Law and varied primarily by industry, occupation, and experience. Entry into occupations was restricted by rigid educational requirements, and incentive payments were small (although not uncommon) and, according to most observers, ineffectual. Workers and managers had only very weak incentives to innovate and risked sanctions for stepping outside the plan. Education was also tightly regulated, as each year the plan specified the precise number of new entrants for each

[1] Brainerd (1998) studies Russia from 1991 to 1994; Chase (1998) contains estimates for 1984 and 1992 in Czechoslovakia; for the Czech Republic, Vecernik (1995) studies 1988-1994, Flanagan (1998) studies 1988 and 1996, and Munich, Svejnar, and Terrell (1999) analyze retrospective data for 1989 and 1991-1996; Kertesi and Kollo (2002) and Campos and Jolliffe (2003) both analyze Hungarian data from 1986 to 1998; Sabirianova Peter (2003) studies retrospective data for 1985 and 1990, and cross-sections for 1994-2000. Most similar to the long time series we study in this paper is Fleisher and Wang's (2004) retrospective data in China from 1950 to 1994.

field. Consistent with Communist development priorities, the educational system strongly emphasized engineering and vocational training relevant to the industrial sector.[2]

The breakdown of this closed, inflexible regime at the beginning of the 1990s came without warning. While wages in state bureaucracies, and for a time in state-owned enterprises, continued to be prescribed by law, the system was not prepared to deal with changes in corporate governance and the sudden growth of new small enterprises. In the old firms, where explicit regulation was replaced by tax-based wage (incomes) policies, there may have been some inertial tendency to stick with the wage grid, which was still officially promulgated. Yet the gradual accumulation of effects from privatization and liberalization likely increased the pressure for firms to rationalize their wage structures.[3] Meanwhile, the Romanian educational system also underwent big changes, as higher education was liberalized and enrollments dramatically increased (Sapatoru, 2001). Concurrently with the shift of employment towards trade and consumer services, students increasingly shifted from technical fields towards humanities, social sciences, and business; and curricula were restructured under the influence of market pressures and international norms.

This context suggests a set of contrasting hypotheses about the changing wage structure in Romania. On the one hand, the tendency for central planners to undervalue education and to compress wage differentials suggests that any earnings premia associated with formal schooling would be small under the socialist regime, and they are likely to increase during transition as the economy liberalizes and moves to a market equilibrium. Furthermore, the usefulness of skills acquired through schooling might rise during transition because of skill-biased shifts in labor demand, improvements in the quality of education, or increases in the "value of the ability to deal with disequilibria" (Schultz, 1975). The opening of international borders, particularly to the West, could increase pressure on the educational premium as more educated workers emigrate to exploit the higher returns on international markets.

On the other hand, expanded access to schooling may have led to a skill-biased relative supply shift, implying a decreased measured return. Moreover, the pre-reform educational system was designed to further the industrialization priorities of the Communist elite, and the

[2] Kornai (1992) contains a general overview of these aspects of socialist economies, while Ben-Ner and Montias (1991) provide some specific discussion on Romania.

[3] Earle (1994), Earle and Oprescu (1995), Earle and Pauna (1996), and Pauna and Pauna (1999) describe Romanian labor markets in transition, while Earle and Sapatoru (1993, 1994) and Earle and Telegdy (1998, 2002) analyze Romanian privatization policies.

value of such schooling might therefore decline in a new market environment.[4] The disruptions of transition might result in a declining, rather than improving, quality of education, reducing the return to recent schooling as well. Finally, the large sectoral shifts associated with an economy-wide restructuring process could imply that the return to schooling is influenced by compositional effects – in either direction.

Theoretical considerations alone, therefore, do not provide a single prediction of the evolution of schooling differentials across the socialist and transition periods. In addition to providing empirical estimates of these differentials from 40 years before to 10 years after transition began, our empirical analysis in this paper exploits information on the nature of Romanian reforms and institutions to try to sort out some of the relevant explanations for the patterns we observe. An examination of the timing of the changes in schooling returns vis-a-vis the timing of liberalization helps assess the plausibility of "movement towards equilibrium." Evidence on changes in the quantity of workers with more and less education is useful to assess the demand and supply interpretations. Some information on the importance of pressures arising from international border opening can be obtained by examining regional and ethnic differences in the schooling premium. Concerning "dealing with disequilibria," we separately estimate the impact of schooling in the private and self-employment sectors, the loci for entrepreneurial behavior in the economy. The possibility of changing value of the educational system can be approached by permitting the estimated return to vary with the time period in which schooling was acquired. Finally, separate estimates of schooling returns by economic sector can, together with information on sectoral shifts, predict the counterfactual return in the absence of the shifts.[5]

In the next section, we describe our data sources, sample composition, and variables. Section 3 contains estimates of the basic earnings functions, while Sections 4 and 5 provide evidence on possible explanations of the observed patterns, the former focusing on relative supply shifts and movement towards equilibrium and the latter on factors that may have shifted relative demand. The final section gives a brief conclusion.

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[4] Flanagan (1998) and Filer, Jurajda, and Planovsky (1999) make this point with respect to the Czech Republic. Kertesi and Kollo (2002) argue more generally that skill obsolescence is an important factor in Hungary.

[5] We follow the previous literature in referring to the coefficient on years of schooling in a conventional earnings function as the "return to schooling," although consideration of issues such as the costs of schooling (monetary and psychic), the measurement of the value of a job (i.e., including fringe benefits and other work conditions), and the problems of estimation (for example, selection bias in schooling decisions) suggest that "wage differential associated with schooling" would be more cautious and apt, although also clumsier.

2. Data

The source of our data is the Integrated Household Survey (IHS) of the Romanian National Commission for Statistics (renamed as National Institute of Statistics since 2001). For the socialist years (back to 1950) and early 1990s, our information is based on retrospective information in the 1994 survey, while for 1994 onwards we use the annual household survey. The IHS started in April 1994, running for 12 months over a changing sample (thus, when we refer to “1994 sample,” this means April 1994 to March 1995). Subsequent years were organized on a similar pattern up to 1997, when the IHS started in April and ended in November. For the rest of the cross sections (i.e., 1998-2000), the IHS started in January and ran for 12 months. Unfortunately, although originally designed as a panel, the data do not permit linking of individual observations across years.

The sample sizes in these data are larger than in most studies of socialist and transition economies. The number of observations available for analysis varies across the cross sections, starting at 25,565 in 1994, falling to 15,508 in 1997, increasing to 21,518 in 1998, and decreasing again afterwards to 17,480 in 2000. Given the relatively small number of yearly observations before 1994, we aggregate these observations into five 5-year periods (1950-1989) and one 4-year period (1990-1993), the latter capturing the initial years of reforms.

A notable change in the survey across the years for the purpose of this paper is that years of schooling are reported directly by respondents only in 1994 and 1995. In order to estimate the return to years of schooling in 1996-2000, we had to impute years associated with educational attainment, a frequent procedure in such data sets. Our method was to compute the median years of schooling for each attainment category in 1994, and then to associate these medians with the corresponding attainment categories in 1996-2000.

Table 1 presents some descriptive statistics on the main variables. Throughout the paper, we provide results for the sample of all employees aged 15-59, but we have also analyzed other age ranges (18-59 and 30-50) and separated the sample by gender, obtaining similar results. The net monthly wage is computed as earnings on the primary job in the previous month minus taxes and other mandatory contributions. The wage variable refers to the previous month in 1994-2000 and the starting wage for jobs held during 1950-1993. Questions may be raised about recall bias in the retrospective information, but it should be borne in mind that starting wages on new jobs are relative easily recalled, particularly in the socialist context of strong stability in wages, prices, and employment. Age bias may also be present, as workers observed to be starting jobs in earlier years tend to be systematically younger than those starting later. All our regressions control for age (experience) in quadratic form, and we have also investigated quartics with similar results, but these problems might still represent significant limitations if we intended a very extensive analysis of the retrospective information. Most of our analysis in this paper concerns the evolution of the differential from 1994 onwards, however, and we use the socialist period data primarily to establish a baseline for the subsequent changes. Moreover, our findings show little fluctuation in the estimated relationships over the entire socialist period, which is inconsistent with large roles played by recall and age bias. To allay any residual concern, we provide estimates of the basic functions using least absolute deviations (LAD) in addition to ordinary least squares (OLS).

Table 1: Summary Statistics, by Time Period

The stability of wages from 1950 to 1989 is clearly shown in the computations of the mean wage, which evolved slowly until jumping up abruptly in 1990-1993, when prices and wages were quickly liberalized. Consistent with aggregate inflation statistics, the mean wage increases rapidly through most of the 1990s. These are nominal wages, but as the cost of living also rose in these years, the average real wage certainly fell. Our concern in this paper is wage differentials rather than the overall level of real wages, so in principle our approach of estimating repeated cross-sections would seem to require no deflation of the dependent variable. The significant inflation during the 1990s, however, requires some within-survey-period adjustments. In each of the retrospective periods (1950-1993), where there are fewer degrees of freedom (as shown in the sample sizes at the bottom of Table 1), we include a quadratic monthly time trend in the equation. In each yearly regression from 1994 to 2000, we include a set of monthly fixed effects.

The sample characteristics in Table 1 also show an average years of schooling at 9.04 in the early 1950s, falling to 8.39 in the early 1960s (possibly due to the Communist regime's active campaign against the intelligentsia), and then rising steadily thereafter, with some acceleration after 1990, to 12.19 in 2000. The increase in years of schooling reflects both the increase in the mandatory education during the communist years and the expansion of educational opportunities in post-socialist Romania. Given the characteristics of the retrospective data, it is not surprising that the potential years of experience tend to be low during the socialist period and until 1994, while afterward the analyzed employees have on average around 20 years of work experience.

Table 2 presents some descriptive statistics for other variables we analyze in the 1994-2000 period. The regional and ethnic distributions are relevant for the possibility that opportunities for emigration have increased the schooling premium. Region is defined by classifying counties (*judete*) on the basis of the "development regions" of the National Commission for Statistics (2000, p. 601), while the ethnic variables reflect the information available in the survey; the means show only minor fluctuations from year to year. The share of employees who have graduated after 1992, which we take as a proxy of post-communist schooling (the variable *NEW*), increased from 0.02 in 1994 to 0.15 in 2000. The figures also show some inter-industry shifts, particularly into service sectors; the figures for agriculture are much lower than from the Labor Force Survey or other official sources, probably because we exclude the self-employed. The biggest shifts concern firm ownership, where the public share falls from 0.86 in 1994 to 0.40 in 2000, the mixed rises from 0.02 to 0.10, and the private from 0.10 to 0.42. These changes reflect the privatization of the Romanian economy, which if somewhat slower than in some neighboring countries, nonetheless changed dramatically during this period.

Table 2: Summary Statistics for New Education, Ownership, Sector, Region, and Ethnicity, 1994-2000

3. Estimating Earnings Functions in Romania, 1950-2000

The basic earnings function we estimate in this paper is the standard relationship due to Mincer (1974):

$$\ln(W) = \beta_0 + \beta_1 S + \beta_2 X + \beta_3 X^2 + \beta_4 F + \sum_t D_t + u, \qquad (1)$$

where the variables are defined in Table 1, the D_t parameterize time to control for general inflation (quadratic monthly time trends in 1950-1993, monthly dummies in 1994-2000), the βs are parameters to be estimated, and u is an error term. Because of some concern about possible measurement error, as discussed above, we estimate using least absolute deviations (LAD) or median regression, as well as by ordinary least squares (OLS).[6] The results from these estimates for cross-sections of employees from 1950 to 2000 are provided in Table 3.

Table 3: Basic Earnings Functions, by Time Period and Estimation Method

Under both estimation methods, we find a small but statistically significant impact of schooling under central planning: our estimates show a fairly constant 3-4 percent premium associated with an additional year of schooling from 1950 through 1989. The slightly higher coefficients in the 1960s might be associated with the industrialization drive that really took off in this period and that relied on wage differentials to induce worker mobility in directions desired by the planners; or they might reflect a recognition that the repression of the intelligentsia during the 1950s had been counterproductive. At any rate, these movements are very slight compared to those beginning in the early 1990s, when the estimated coefficient begins to trend upward steadily, more than doubling – to 8.5 percent – by the year 2000. Throughout the retrospective data analysis, the LAD coefficients are much smoother than the OLS, and in particular they show a smaller jump in 1990-1993, but from 1994 on there is little to choose between them. The results provide new evidence, based on longer time series than previously available, concerning the low "return to schooling" under socialism and the dramatic rise in the return during transition.[7]

Although not the focus of this paper, the results for the other variables are also interesting. The return to the first year of experience rises in the 1990s compared to the pre-reform period. The concavity of the experience profile also tends to increase, consistent with results in other countries. Finally, the coefficient on the female dummy is consistently negative, and the magnitude tends to be larger in absolute value in the transition period.[8]

Our findings for the schooling coefficient may be compared with those obtained in other studies of transition economies. As we noted above, our paper provides the first analysis of the evolution of the wage impact of schooling in Romania from the socialist to the transition period. Skoufias (2003), however, provides estimates for 1994, and our results for that year are very

[6] As a further check on the influence of possible measurement error, we also estimated the earnings functions with samples that excluded that top one percent of earners in each period; the results were similar to those reported here.

[7] Motivated by the possibility that participation rates of low earners might be falling over this period, particularly those of younger people (who might stay in school), older people (who might retire early), and women (who might be withdrawing to care for children), we also estimated all equations for the central age-group of 30–50 years old, and for men and women separately. The qualitative patterns in these results are again very similar to those reported here. We also discuss changes in participation patterns by schooling category below.

similar to his.[9] Concerning studies of other economies that examine the evolution over time, Chase (1998) finds a much smaller return during the socialist period in Czechoslovakia but a similar figure for 1993 to ours for 1994. Munich, Svejnar, and Terrell (1999) also report a lower return before transition in the Czech Republic, while their estimate of 5.8 percent in 1996 is similar to Flanagan's (1998), and both are slightly smaller than our 6.7 percent estimate for Romania. For Russia from 1991 to 1994, Brainerd (1998) estimates an increase from 3.1 to 6.7 percent for men and 5.4 to 9.6 percent for women. For Hungary, Campos and Jolliffe (2003) report an estimated return of 6.4 percent already in 1986, rising to 11.2 percent by 1998. Using the same data, Kertesi and Kollo (2002) report that the return to education in Hungary rose quickly from 1989 to 1992 but then leveled off. Our findings differ in showing a steadier and more gradual evolution of the estimated return in Romania. In Fleisher, Sabirianova Peter, and Wang (2004)'s summary of estimates of schooling returns across a number of transition economies, the mean estimate is about 4 percent in the late 1980s, rising to 8.8 percent in 2000; our estimates for Romania are very close to these.

4. Supply, Demand, and Movement toward Equilibrium

As this discussion makes plain, the pattern of increasing wage differentials associated with schooling has been well-documented in a number of transition countries, and our results so far provide evidence of a similar pattern in Romania. But what factors might explain these dramatic changes? Although data limitations prevent us from a detailed investigation of all the possibilities, we are able to provide some evidence relevant to a number of hypotheses. A first group of these concerns basic supply and demand analysis: an increase in relative pay associated with longer schooling may reflect an adjustment to equilibrium wage relativities, it could be due to a contraction in the supply of more educated workers, or it could reflect skill-biased shifts in labor demand. In this section, we consider these broad categories of explanation, before moving on in the next section to the specific factors that may underlie relative demand shifts.

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[8] The widening gender gap in our data is an exception to Brainerd's (2000) analysis of gender differentials in several East European countries (not including Romania), but it is consistent with her finding for Russia and Ukraine. Why Romania should be an exception to the East European pattern is a topic worth further research.

[9] Skoufias (2003) measures schooling as a set of dummies for educational attainment rather than years of schooling, and his sample differs in several ways (maximum age of 65, restriction to individuals interviewed in 1994), but we receive results similar to his when we estimate using attainment dummies with our sample.

The first group of hypotheses can be illustrated with a simple demand-supply diagram, as in Figure 1. The horizontal axis measures the average schooling in a population while the vertical measures the marginal return to additional schooling. The demand and supply functions are expressed in relative terms, the former showing how relative earnings vary with average schooling, and the latter measuring the willingness of workers to acquire additional schooling if faced with a higher return. We have drawn the supply function as relatively inelastic due to the presumed time lags for workers responding to different incentives.

Figure 1: Understanding Changes in Relative Wage ($\partial W/\partial S$) and Quantity of Schooling (S)

Three hypothetical situations are portrayed: a below equilibrium level of the relative wage at the very beginning of transition, labelled W_{1990}; the result of moving to equilibrium with simultaneous outward shift of both demand and supply in the middle of the transition process, W_{1995}; and the result of further outward shifts, W_{2000}. W^*_{1990} refers to the relative wage in 1990 if workers had been paid their marginal products.[10] The relative importance of adjustment to equilibrium at the beginning of transition can be measured in the diagram as $(W^*_{1990}-W_{1990})/(W_{1995}-W_{1990})$. We approach an analysis of this issue in two ways: first, we consider the temporal pattern of the growth in the schooling coefficient in relation to the liberalization of labor markets in Romania; second, motivated by the possibility of inertial wage setting practices for tenured workers, we study the evolution of returns to schooling by cohort.

The first type of analysis comes directly from the figures in Tables 1 and 3, and we have provided a graphical analysis in Figure 2 (using the LAD coefficients from Table 3). The liberalization and adjustment hypothesis would imply a sharp jump in the return to schooling around the time of the dramatic policy changes of the early 1990s, followed by a fairly constant return in the later years.[11] Instead, the figure depicts continuous increases throughout the 1990s, with only a small share of the adjustment taking place in any particular year. The schooling coefficient does jump more in the early 1990s than later on, but the continuing upward trend would seem to provide *prima facie* evidence directly contradicting the hypothesis.

Figure 2: Observed Changes in Relative Wage ($\partial W/\partial S$) and Quantity of Schooling (S)

[10] The definition of productivity during the socialist period is somewhat problematic, as the system had different goals, prices, and wages; to avoid this confusion we refer to the 1990 situation, when the goals of a market economy are assumed, yet wages were still controlled.

[11] See Earle and Oprescu (1995) for a discussion of wage regulations and policies in the early 1990s. The biggest change came in February 1991, when wage setting was liberalized, although some controls continued to be imposed in the state sector. Below, we analyze differences in the schooling wage premium by ownership type.

Perhaps this view is too strict, however, as it is likely that individual workers' wages may respond sluggishly and institutional factors may intervene, particularly in the short run, so that the adjustment toward equilibrium takes place only gradually. In this case, however, it would imply that the greatest adjustments would be on the margin: for instance, younger cohorts of workers and those just hired. For this reason, we also estimate a modified version of equation (1):

$$\ln(W) = \beta_0 + \beta_{10}S + \beta_{11}XS + \beta_{12}X^2S + \beta_2X + \beta_3X^2 + \beta_4F + \textstyle\sum_t D_t + u \,, \tag{2}$$

which permits the β_1 coefficient on S in equation (1) to vary with work experience. We pool the years 1970-1989 together for this analysis and also estimate on the 1990-1993 time period and for each year thereafter. The results for $\partial W/\partial S$, graphically displayed in Figure 3, show that initially the schooling wage premium rises more for younger cohorts (β_{10} is larger) and declines with experience (β_{12} is larger in absolute value), but in fact the estimated return grows rather steadily for each experience group. By the late 1990s, the profile has nearly converged to a profile that is a simple 4 percentage point upward shift of the socialist profile, with little difference in shape.[12]

Figure 3: Evolution of the Experience Profile of Returns to Schooling

The data, therefore, provide only a little support for the simple "movement to equilibrium" interpretation. In terms of Figure 1, W^*_{1990} appears to differ relatively little from W_{1990}, at least compared with the shifts implied by the magnitudes of W_{1995}-W_{1990} and W_{2000}-W_{1995}. Most of these increases must instead be explained by shifts of the relative supply or relative demand functions.

The possibility that a contraction of relative supply caused the rising measured return to schooling is directly contradicted by the increased level of education in the Romanian population. As demonstrated by Figure 2, which portrays the evolution of the average years of schooling and the estimated wage premium associated with schooling over the period 1970 to 2000, the relative supply of educated workers expanded steadily. The supply-side changes, therefore, would have served to reduce, not increase, schooling returns. The continual

[12] Similarly motivated by the possibility of inertial wage setting for incumbent workers coupled with greater adjustment on the margin (i.e., for those recently hired), we estimated similar equations for 1994 and 1995 with two alternative measures of recent hiring – hired in the previous year and hired since 1991 – based on the job tenure variable, which is available for those two years only. The results were consistent with this motivation, implying a 1-2 percent greater schooling premium for the recently hired, but this small difference (coupled with the small fraction of recently hired workers) is insufficient to account for more than a negligible amount of the growth in the schooling coefficient over this period.

movement up and to the right in Figure 2 appears to be tracing out the equilibria shown in Figure 1.[13]

The relative expansion of skilled worker supply took place at the same time as the liberalization of Romania's borders opened up the possibility of emigration, which may have been especially attractive for better educated workers who could exploit the higher schooling premium to the West. Although such emigration clearly did not offset the overall relative supply increase within Romania, it is possible that the pressure raised the schooling premium in certain groups, those with the greatest tendency to emigrate. Suppose, for example, that the relative supply elasticity is identical for all groups but that the relative demand function faced by groups prone to emigrate happens to be less elastic. In this case, a leftward shift in relative supply of the emigration-prone could raise the schooling premium overall, even if the rightward shift of the non-emigration-prone was great enough to simultaneously raise the overall average level of schooling. Two simple tests of this argument involve two different ways of proxying the tendency to emigrate, the first based on region (distance from the Western border) and the second based on ethnicity – Hungarians and Germans, who have enjoyed not only valuable language abilities but also preferred emigration status in Hungary and Germany, respectively. In both cases, we rely upon variants of equation (1) involving interactions between schooling and the relevant variables: region in the first case and ethnicity in the second.

Summary statistics for the region and ethnicity variables were shown in Table 2, while the results of the regression analyses appear in Tables 4 and 5. Concerning variation in the estimated schooling return by region, shown in Table 4, the coefficients of interest involve the interactions between schooling and the western regions – Southwest, West, and Northwest – which are located closest to Hungary and job opportunities in the European Union and thus may be expected to have the highest returns. Contrary to this hypothesis, all these coefficients are negative, and occasionally they are even statistically significant at conventional levels. Concerning variation by ethnicity, the coefficients of interest are the interactions of schooling with Hungarian and German background, and again the results are inconsistent with the hypothesis that an improvement in the relative opportunities for more educated workers in these ethnic groups has effectively shifted their relative supply functions backwards.

[13] We have also examined the evolution of employment-population ratios for three educational groups (S<12, S=12, and S>12) and find some tendency for the employment probability to decline more for the less educated compared with the more educated group. Thus, the rise in average educational attainment is higher among employed individuals than in the population as a whole.

Table 4: Variation in the Return to Schooling by Region

Table 5: Variation in the Return to Schooling by Ethnicity

Overall, therefore, we find no evidence of any role for supply shifts in explaining the rapidly rising return to schooling in Romania. Indeed, the large supply shifts we observe would imply a decline, not an increase, in the schooling effect. Furthermore, the increases in student enrollments and average worker education imply a still greater reduction in the wage differential associated with schooling as long as schooling and ability are correlated, for the expanded opportunities for schooling would result in lower average ability at higher levels of schooling, lowering the schooling coefficient over this period. The shifts in relative demand must have been large enough to offset these negative effects from the supply side as well as to account for the large observed rises in both the quantity and price of educated labor.

5. Explanations: Relative Demand Shift Factors

The evidence so far clearly suggests that the rising return to schooling in Romania during the 1990s must be explained by large outward shifts in the relative demand for more educated workers. What factors could have led to the increased relative productivity of more educated workers that would underlie such shifts? A first possibility is an increase in the quality of education. Second, demand could shift due to skill-biased technical change. Third, even if there was little change inside Romania, it is possible that international opening of the economy could effectively raise relative demand, putting upward pressure on skill differentials to bring them in line with neighboring countries. Fourth, product demand shifts across industries – using different technologies and therefore providing different rewards for schooling – could produce compositional effects in the changes in the estimated schooling coefficients. Fifth, similar compositional effects could occur due to shifts across ownership forms, in particular from the state to the private sector, where wage-setting mechanisms are likely to differ significantly. Finally, the opportunities for entrepreneurship in the unstable environment of transition could increase returns if education is associated with a greater ability to "deal with disequilibria." We consider each of these possible explanations in turn.

The first possibility, improvements in the educational system, can be thought of as technological changes to the human capital production function. This idea has only recently surfaced in discussions of rising skill differentials in the West (Bowlus and Robinson, 2004), but it has been more common in East European discussions of these issues (e.g., Kertesi and Kollo, 2002). A popular view among educators in the region is that the educational system has become less productive, the strenuous standards of the socialist system – particularly in mathematics and technical fields – having deteriorated under the lax discipline of transition. If true, this would imply a decreased return to schooling, *ceteris paribus*. As a crude test of these possible changes in the educational production function, we distinguish workers who graduated after 1992 as having "new education." The means by year for this variable (*NEW*) are shown in Table 2.

Our method is to interact *NEW* with *S* in another extension of equation (1). The results are shown in Table 6, and they indicate a small premium for post-communist schooling in 1994 and 1995 of about 2 percentage points. The estimated coefficient shrinks to 1 percent and becomes statistically insignificant in 1996, however, and thereafter is completely negligible in size as well as statistically insignificant. It is also noteworthy that those with new education receive sharply lower earnings (i.e., a lower intercept) in 1994 and 1995, but this difference converges towards zero over the 1990s. In any case, new entrants are clearly not particularly highly rewarded in the Romanian labor market during this period, and the evidence does not appear to support the hypothesis that improved education has raised the productivity differential associated with more schooling.

Table 6: Variation in the Return to Schooling by New versus Old Education

A second possible explanation for the outward relative demand shift could be skill-biased technical change. The notion that advances in information technology account for increased wage inequality has been extremely fashionable in the U.S., but unfortunately it is very difficult to measure. In our data, there is no variable to proxy for computer usage or technology adoption by the firm. Common sense, however, suggests that it is implausible that technology change, at least of the conventional sort, is a major factor. For one thing, the increase in the wage impact of schooling is much faster in Romania during the 1990s than in Western economies in the entire second half of the twentieth century. Indeed, as noted by Card and DiNardo (2002), the increase in the schooling premium in the U.S. had taken place by 1990, with little change thereafter. Even if Romania started transition in a technologically backward state, investment was very low through most of this period, so adoption of new technology was probably similarly sluggish.[14] Some direct evidence from firm surveys appears in Commander and Kollo (2004) and Earle, Pagano, and Lesi (forthcoming); both studies show low levels of information technology usage, and the former shows that adoption is largely uncorrelated with the rise in the skill premium in the sampled firms. Perhaps technological change in a broader sense including not only physical machinery but also new types of organizational practices might be responsible, although these are even harder to measure.[15] We return to a discussion of such changes below.

Another broad category of explanation concerns changes in the composition of the Romanian economy. Research on the increasing schooling premium in the U.S. associates sectoral shifts with changes in product demand, and similarly we may consider the rise of the service sector and the decline of heavy industry in Romania as reflecting the substitution of consumer preferences for central planning in the determination of product demand. For current purposes, we consider shifts across a crude division of the economy into 3 sectors: agriculture, industry, and services.[16] The main hypothesis of interest is that the return to schooling is higher in the services sector (due, for example, to different technology), so that a rise in services leads to a composition of the economy with a higher weight on the wage differential in services.[17] We again employ an interactions specification, with the results shown in Table 7.

Table 7: Variation in the Return to Schooling by Sector

Industry is the omitted category; thus the coefficient on *S* measures $\partial W/\partial S$ in the industrial sector, while the coefficients on the interaction terms show the difference between the return in agriculture or services from that in industry. The estimates imply an approximate 1 percent additional premium for schooling in services compared with industry, but this difference is small and falls somewhat over these years. Moreover, the level and growth in the estimated return to schooling in industry look similar to those for the whole economy. These results provide little support for a major role of sectoral shifts in explaining the rising wage premium.[18]

[14] The share of investment in GDP, calculated from official figures in National Commission for Statistics (various issues), was 29.6 percent in 1989, 14.2 in 1991, 16.1 in 1994, 16.3 in 1998, and 10.7 in 2000.

[15] Brynjolfsson and Hitt (2000) argue that the main effect of computerization works through the complementary organizational changes accompanying new technology adoption; in this case, however, the skill-bias effect should still be correlated with technical change. This implies that other types of organizational change may be more significant.

[16] See Earle (1997) for a more detailed discussion of interindustry mobility of workers in Romania.

[17] The sectoral shares of employees in our data differ from those in official statistics because of large numbers of self-employed in both agriculture and services. If we include self-employed, the shares of agriculture, industry, and services in 2000 would be 36 percent, 25 percent, and 39 percent.

[18] A similar analysis with 15 disaggregated industries also finds no indication that interindustry shifts in employment could contribute significantly to the rise in the coefficient overall.

Another variety of compositional shift concerns ownership types. As shown in Table 2, the Romanian economy underwent dramatic changes by ownership during the 1994-2000 period, with a substantial decline in the fraction of workers reporting their employer was state-owned (from 86 to 40 percent), and corresponding rises in the fraction private (from 10 to 42 percent), mixed (from 1.8 to 10.5 percent), and an unknown "other" (from 0.2 to 5.5 percent). Our motivation for studying these ownership forms is the possibility that they differ in organizational practices, due to legal regulations, firm objectives, or corporate governance. These practices may result in deviation of relative wages from relative productivity ratios of workers within a firm. The specific hypothesis is that private firms – placing a higher weight on profits, feeling more pressure from market competition, and facing harder budget constraints – are less likely to provide such rents to low-skilled workers than the state sector. We provide evidence on this hypothesis with a test analogous to those above, namely by adding to equation (1) interactions of ownership type with schooling. State ownership is the omitted category.

The results, presented in Table 8, imply a statistically significantly higher schooling wage premium in privately owned firms. Interestingly, the estimated magnitude follows a roughly inverted-U trajectory, rising from 1994 to 1996 and falling thereafter. This difference in wage-setting behavior in the private sector, combined with the rising private share in total employment, may partially account for the overall growth in the aggregate schooling return. The contribution is not large, however: the private sector added about 0.2 percentage points to the aggregate return in 1994 and about 0.7 in 2000. Meanwhile, the estimated return in the state sector grows by 2 percentage points (from 5.7 to 7.7, as shown in the table).

Table 8: Variation in the Return to Schooling by Ownership of Employer

Our findings suggest that, contrary to a number of hypotheses, the rise in the wage premium for additional schooling was both gradual and broadly based. It was not concentrated in only some sectors of the Romanian economy but affected all sectors without many differences among them. The fact that the private sector appears to have led the increasing trend is suggestive, however, as it implies that changes in organizational practices may be part of the story.

What sorts of organizational practices could be relevant, and what changes in the economic environment could have brought them about? One possibility is raised by recent research on skill differentials in the U.S., which maintains that the effect of technological change works through organizational practices to raise the relative productivity of more skilled workers. Brynjolfsson and Hitt (2000), for instance, point to the ways computers have enabled practices such as flexibility in equipment design and job assignments, lower levels of inventories, more outsourcing, more participation in decision making, and flatter hierarchies. But we have argued that the rising skill differential in Romania (and other transition economies) took place much more quickly than can be explained by investments in new technologies and the even slower adoption of such practices.

Our final hypothesis, therefore, concerns a different set of practices that involve particular types of skills and tasks: finding creative solutions to problems, recognizing and exploiting new opportunities, innovating rather than simply following orders. The socialist system provided workers and managers with few incentives to display individual initiative and exercise these qualities.[19] Not only were many economic decisions prescribed by the plan, the stability of the system meant that there were few gains from searching for new opportunities; innovation and exceeding the plan targets could even be penalized, for instance through the "ratchet effect." In the transition, however, the abilities to think "outside the box" and to act entrepreneurially became extremely important, probably even more so than in stable market economies. If education increases these abilities to "deal with disequilibria," as argued by Schultz (1975), then the relative productivity of workers with more schooling will rise.

The problem is how to measure or provide some evidence on this effect. We do so indirectly, by analyzing the returns to schooling among the self-employed. For this purpose, we consider the nonagricultural self-employed as entrepreneurs, as they are typically treated in the literature on this topic.[20] Comparing with our estimated coefficient for employees, if we find a similar or lower schooling coefficient for self-employed, then this would imply a rejection of the argument, while finding a substantially higher coefficient would be consistent with it. The return to schooling among entrepreneurs might be expected to first rise and then fall, as the scope for exploiting new opportunities initially rises (as liberalization increases and the opportunities are revealed) and then declines (as the opportunities are exhausted).

[19] One should not entirely discount the usefulness of creativity in solving such problems as the supply breakdowns endemic under central planning; the assumption here is only that the scope for and return to exercising creative initiative were greatly attenuated compared to a market or transition economy.

[20] See, e.g., Evans and Leighton (1989), Fairlie and Meyer (1996), or Hamilton (2000). Consistent with most literature, we omit the agricultural self-employed from the analysis as they are less likely to be genuine entrepreneurs, particularly in Romania, where the land privatization policy resulted in tiny family farms.

Defining and measuring the income of entrepreneurs is always a difficult problem, but in the case of the IHS a special section of the questionnaire provides unusually detailed and precise information: gross revenue from entrepreneurial activities, capital inputs, material inputs, labor costs, taxes, and in-kind payments – all with respect to the reference month.[21] We define net income as the first variable minus the sum of all the others and use this as the dependent variable in the conventional earnings regression (1). Table 9 presents estimates for the sample of nonagricultural self-employed, aged 15-59, in the IHS from 1994 to 2000.

Table 9: Return to Schooling for the Nonagricultural Self-Employed

The estimated coefficient on *S* is larger for the nonagricultural self-employed than for employees in all years. The coefficient grows strongly until 1998, when it peaks at 15.5 percent, and then declines somewhat thereafter.[22] The pattern is not due to changes in the supply of individuals engaged in self-employment, as the fraction of total employment accounted for by the nonagricultural self-employed steadily expanded, cumulatively nearly doubling (from 3.58 to 6.03 percent) in just six years from 1994 to 2000.[23]

These results are consistent with the proposition that education plays an important role in enhancing the ability of workers to deal with disequilibria. We believe they shed light not only on the self-employed, but also on the increased return to education among employees. Employees may also be involved in entrepreneurial activities, in the sense of recognizing and exploiting new opportunities. If education enhances the ability of the self-employed to act creatively, then it may be inferred that it has a similar effect for employees as well.

[21] In-kind payments, which would mostly refer to crops given to workers, are available only in 1994 and 1995, but they would represent subtractions from gross revenue in later years. The use of data for a reference month is somewhat problematic, but we have little alternative with the data available.

[22] Studies of these relationships in other countries have found varying results: Gill (1988), Borjas and Bronars (1989), Evans and Leighton (1989), and Fairlie and Meyer (1996) find a higher return for self-employed, but Rees and Shah (1986), Earle and Sakova (2000), and Hamilton (2000) find the opposite.

[23] The fraction would of course be much greater if we followed the convention of calculating the rate in nonagricultural employment: the relevant figures for 1994 and 2000 in this case would be 4.7 and 8.3 percent, respectively.

6. Conclusion

This paper makes a number of contributions to research on the growth in the estimated return to schooling during the transition from socialism. Ours is the first paper to examine the changes in the return for Romania, a relatively large country in Eastern Europe that has been somewhat neglected by researchers. Our paper is also one of very few to contain information for long periods during both the central planning and transition years: we analyze 40 years and 11 years of data for the two periods, respectively. Our estimates of basic earnings functions in Romania reinforce previous research findings from other countries that the schooling wage premium was low under central planning (although our point estimates, at around 3-4 percent, are somewhat larger than those for the Czech Republic and smaller than those for Hungary, for instance) and that it grew substantially during the transition years – more than doubling in our analysis of Romanian data through the year 2000.

Our paper also goes beyond estimating the schooling coefficient to assemble evidence concerning a number of explanatory hypotheses for the observed patterns. We first investigate the conventional explanations for an increased schooling premium in Western research, including relative supply shifts, product demand shifts, and skill-biased technical change. The rise in average schooling in our data is inconsistent with an overall contraction in supply of more educated workers in Romania, and the lack of evidence of higher returns for workers in the West and for ethnic groups with better emigration possibilities (Germans and Hungarians) leads us to reject any role for border liberalization in putting upward pressure on the schooling differential. Our analysis of interindustry variation in estimated schooling returns provides no evidence of a significant impact of product demand shifts. The possibility of skill-biased technical change is difficult to measure and cannot be completely discounted, but the much faster pace of increase in the measured schooling return in Romania compared to the West and the very low level of investment during the same period undercut the plausibility of the large or exclusive role assigned to this factor in many studies of Western economies.

We therefore consider a set of additional hypotheses that we derive from a broader understanding of Romania and other transition economies. First among these is the possibility that the increased return reflects a movement from centrally planned determination to an equilibrium in which relative wages more closely reflect relative marginal products. While again we cannot unequivocally reject this hypothesis, which has dominated most previous research on returns to education in transition economies – at least implicitly – the rather gradual pace of the growth in returns throughout the 1990s, even among new cohorts of recently hired workers, provides evidence against a dominant role for this factor.

The results of our analysis also challenge interpretations based on improvements from the socialist to the transition period in the quality or value of formal schooling. We find neither that education received during the socialist period lost value when market reforms were introduced, nor that newly acquired education after 1990 was consistently valued much higher in the labor market. Indeed, while of course the share of the work force with "new education" steadily increased through the 1990s, the estimated return is significantly larger than that for "old education" in 1994 and 1995, and the difference is negligible and statistically insignificant thereafter. Nevertheless, the overall return to education continued to steadily increase.

Our analysis does find support, however, for a category of explanations that has received little attention in the literature: organizational and institutional changes that increase the value of education. The two main pieces of evidence for this hypothesis are the greater wage effects of schooling among private sector employees and among self-employed entrepreneurs, both of which grew substantially in their share of Romanian employment over this period. The differential in the return averages 1.8 percent for the private sector and 5.0 percent for entrepreneurs, and the evolution of both displays a pronounced inverted U-shape over the 1994-2000 period. Our interpretation of these results is that the adoption of new organizational practices, particularly the higher rewards for individual initiative, increased the value of education within the private sector, while the possibilities of exploiting new opportunities did the same even more so among entrepreneurs. The state sector, meanwhile, was itself gradually commercializing, undergoing organizational change, and experiencing increased labor market pressure to conform to the wage differentials in the growing rest of the economy. The inverse U-shape reflects the leadership of private sector and entrepreneurial returns in pushing the more sluggish state sector in this direction, as well as the gradual exhaustion of great opportunities for dealing with the disequilibria of economic transition.

The analysis we have carried out provides support for these interpretations, but the data are insufficient to refute or substantiate them entirely. Therefore, it is appropriate to conclude by reiterating some important caveats about our work. We should again emphasize that our analysis suffers from the standard problems in studies of "returns to schooling" in that we observe only wages, not other economic or psychic benefits from work, we do not observe costs of acquiring education, and we cannot control for self-selection in individual educational choices. The transition context may particularly aggravate the first two of these problems, as fringe benefits and work conditions changed drastically as did individual variation in schooling costs, with the entry of new private educational institutions and the introduction of the practice of charging fees to some students even in state organizations. Concerning the third problem, we may take the acquisition of schooling-based skills under central planning as exogenous to earnings during the transition, particularly under our argument concerning the large increase in the value of the ability to deal with disequilibria. Thus, the transition context may partially ameliorate this standard problem.

We should also emphasize an important caveat about our analysis of earnings functions prior to 1994, which are based on retrospective questions asked of respondents in 1994. As always, questions about the reliability of such data may be raised, and the results should be treated with caution. Indeed, the relatively low R^2 that we obtain in most of the pre-1994 period certainly suggests the possibility of higher measurement error during this period. To avoid mistaken inferences, we estimate our equations on a variety of samples, including eliminating outliers, and we use LAD as well as OLS estimation methods. All the results from these different approaches show great stability in the estimated schooling coefficient over the entire 40 years, which suggests that mistakes in answering the retrospective questions are not leading to systematic biases.

A final caveat concerning measurement problems applies to nearly all the hypotheses we consider for the observed pattern of increasing return to schooling. Lack of information prevents us from undertaking a more thorough analysis of schooling quality, product demand shifts, and technical change, for instance. We do find little evidence supporting major roles for these factors, but further analysis based on better data would certainly be useful. Concerning the evidence we find for our hypothesis that the transition involves an increased value of education in dealing with disequilibria, data limitations again prevent us from measuring important factors such as creativity, innovation, and initiative. Our findings of higher returns to education in paid private sector work and in entrepreneurship cannot be considered decisive, but we find them highly suggestive of the value of education in a disequilibrium period full of opportunities.

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Figure 1: Understanding Changes in Relative Wage (∂W/∂S) and Quantity of Schooling (S)

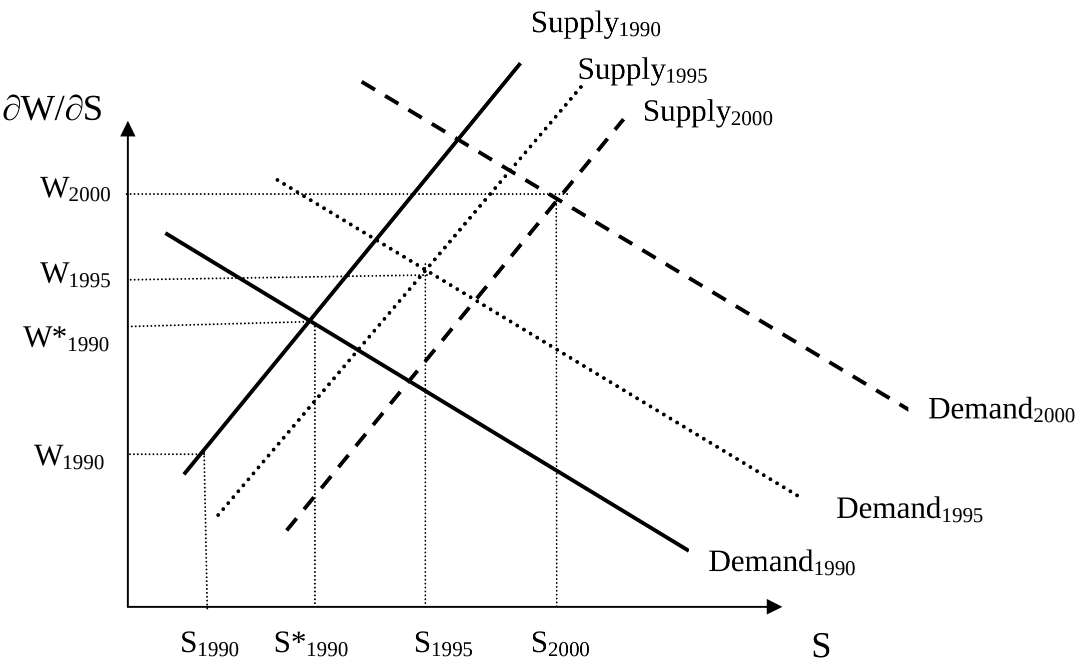

Figure 2: Observed Changes in Relative Wage ($\partial W/\partial S$) and Quantity of Schooling (S)

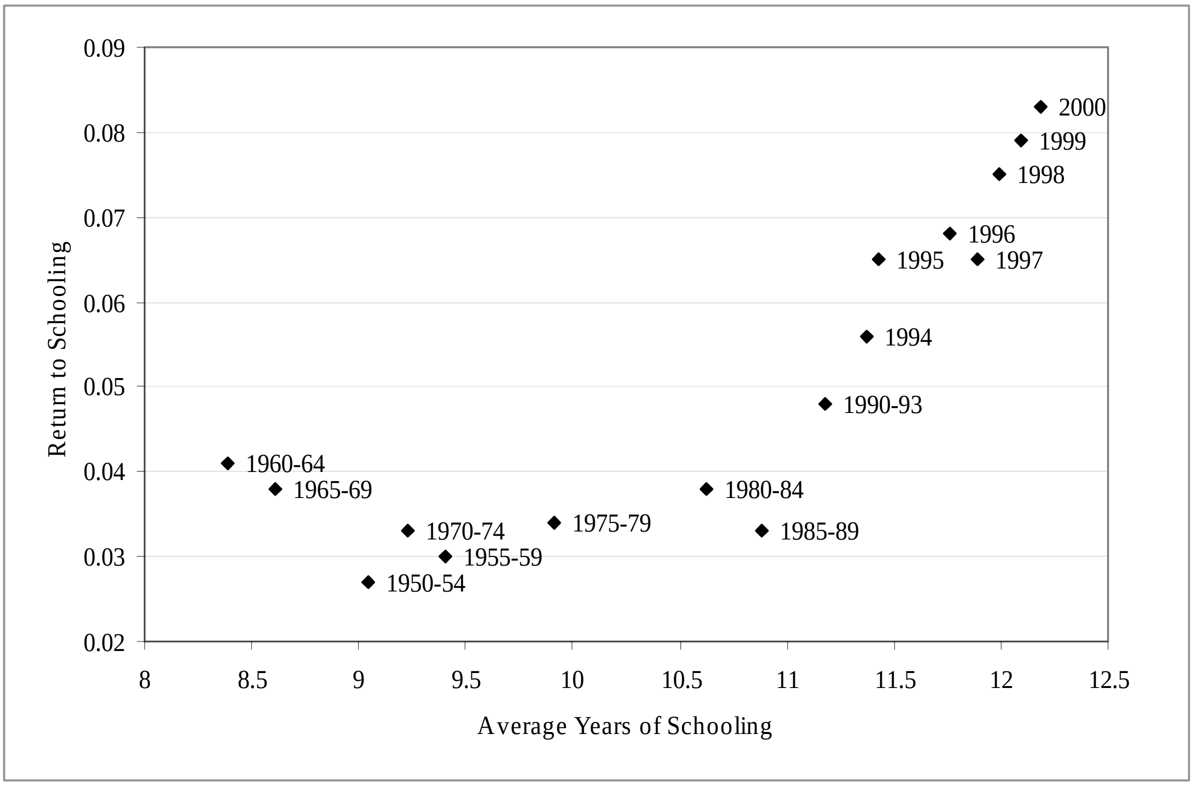

Figure 3: Evolution of the Experience Profile of Returns to Schooling

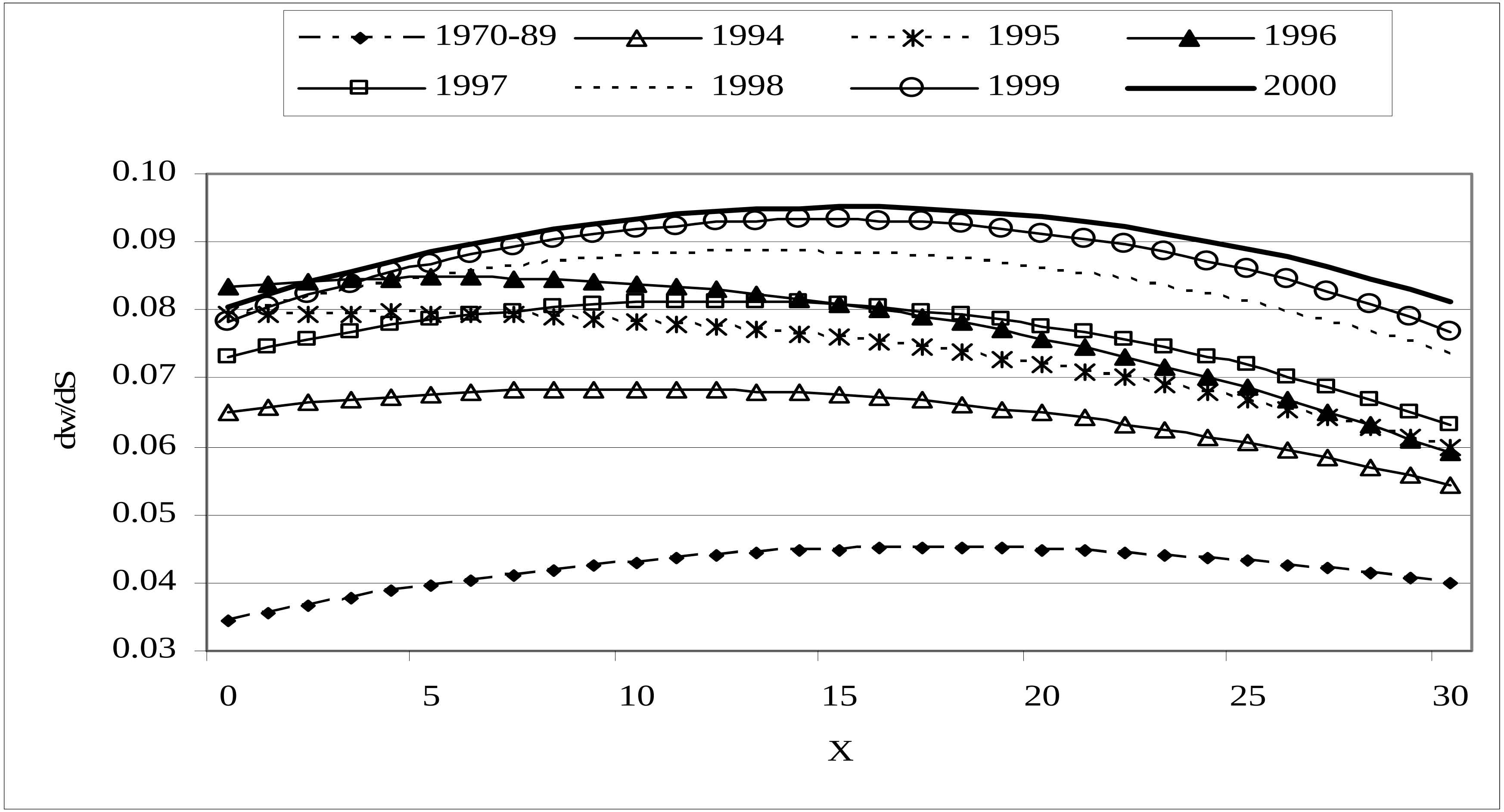

Table 1: Summary Statistics, by Time Period

	1950-54	1955-59	1960-64	1965-69	1970-74	1975-79	1980-84	1985-89	1990-93	1994	1995	1996	1997	1998	1999	2000
W	0.63	0.61	0.76	0.97	1.25	1.71	1.90	2.11	16.00	133.59	192.78	287.35	513.36	857.80	1231.38	1938.38
	(0.96)	(0.34)	(0.47)	(0.77)	(0.67)	(5.00)	(3.04)	(1.55)	(27.27)	(72.82)	(98.71)	(163.38)	(310.14)	(484.83)	(658.47)	(1142.07)
Ln(W)	-0.76	-0.63	-0.46	-0.21	0.11	0.32	0.53	0.65	1.96	4.77	5.15	5.53	6.11	6.63	7.01	7.44
	(0.69)	(0.52)	(0.65)	(0.63)	(0.50)	(0.51)	(0.43)	(0.39)	(1.28)	(0.50)	(0.48)	(0.51)	(0.50)	(0.49)	(0.47)	(0.50)
S	9.04	9.40	8.39	8.61	9.23	9.91	10.62	10.88	11.18	11.37	11.43	11.76	11.89	11.99	12.09	12.19
	(3.70)	(3.88)	(3.62)	(3.52)	(3.29)	(3.29)	(3.21)	(2.79)	(2.49)	(2.87)	(2.78)	(2.60)	(2.56)	(2.52)	(2.50)	(2.41)
X	4.48	6.35	10.88	12.52	12.50	12.34	11.09	9.95	8.91	20.26	20.21	19.81	20.10	20.14	20.17	20.08
	(4.22)	(5.51)	(6.70)	(8.41)	(10.05)	(11.11)	(11.19)	(11.19)	(9.55)	(10.42)	(10.30)	(10.15)	(10.04)	(9.95)	(9.83)	(9.74)
F	0.27	0.38	0.33	0.35	0.42	0.46	0.43	0.46	0.43	0.41	0.41	0.43	0.43	0.44	0.45	0.45
N	459	609	854	805	1237	1339	1676	2606	1228	25565	23644	23919	15508	21518	18963	17486

Note: *W* is net monthly wage (thousand Romanian lei), *ln(W)* is the natural log of *W*, *S* is schooling (years), *X* is potential experience (years), *F* is female dummy, and N is the number of observations. Standard deviations are shown in parentheses (for continuous variables).

Table 2: Summary Statistics for New Education, Ownership, Sector, Region, and Ethnicity, 1994-2000

	Definition	1994	1995	1996	1997	1998	1999	2000
Region								
BUCHAREST		0.127	0.112	0.113	0.116	0.108	0.112	0.113
NORTH-EAST		0.134	0.137	0.133	0.132	0.135	0.130	0.132
SOUTH-EAST		0.123	0.123	0.121	0.122	0.126	0.126	0.114
SOUTH		0.153	0.151	0.155	0.148	0.149	0.146	0.145
SOUTH-WEST		0.105	0.106	0.103	0.104	0.106	0.107	0.114
WEST		0.093	0.096	0.096	0.094	0.098	0.093	0.096
NORTH-WEST		0.135	0.144	0.139	0.144	0.141	0.145	0.152
CENTER		0.130	0.132	0.142	0.140	0.137	0.140	0.133
Ethnicity								
ROMANIAN		0.922	0.919	0.915	0.916	0.916	0.917	0.912
HUNGARIAN		0.063	0.068	0.070	0.069	0.068	0.069	0.074
GERMAN		0.003	0.002	0.003	0.003	0.002	0.003	0.002
ROMA		0.005	0.004	0.006	0.006	0.008	0.006	0.006
OTHER		0.007	0.006	0.006	0.006	0.006	0.005	0.006
New education								
NEW	graduated after 1992	0.019	0.034	0.059	0.078	0.098	0.122	0.147
Ownership type								
STATE	state	0.864	0.806	0.753	0.705	0.622	0.481	0.404
PRIVATE	private	0.100	0.149	0.184	0.224	0.268	0.343	0.423
MIXED	mixed	0.018	0.029	0.048	0.057	0.094	0.119	0.105
COOP	cooperative	0.016	0.014	0.012	0.012	0.013	0.013	0.011
OTHER	other ownership	0.002	0.002	0.003	0.002	0.002	0.042	0.055
Sector of employer								
INDUSTRY	industry	0.446	0.432	0.441	0.436	0.419	0.408	0.410
AGRIC	agriculture	0.085	0.085	0.074	0.068	0.063	0.057	0.047
SERVICES	services	0.469	0.483	0.485	0.496	0.517	0.535	0.543

Note: Regions are defined on the basis of National Commission for Statistics (2000, p. 601).

Table 3: Basic Earnings Functions, by Time Period and Estimation Method

	1950-54	1955-59	1960-64	1965-69	1970-74	1975-79	1980-84	1985-89	1990-93	1994	1995	1996	1997	1998	1999	2000
OLS																
S	0.031	0.024	0.047	0.046	0.039	0.042	0.043	0.034	0.064	0.059	0.067	0.067	0.069	0.078	0.082	0.085
	(0.015)	(0.008)	(0.008)	(0.008)	(0.006)	(0.005)	(0.004)	(0.003)	(0.011)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
X	-0.003	0.016	0.006	0.014	0.020	0.009	0.013	0.015	0.036	0.022	0.024	0.027	0.032	0.027	0.026	0.031
	(0.022)	(0.010)	(0.010)	(0.010)	(0.004)	(0.004)	(0.003)	(0.002)	(0.008)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
$X^2/100$	0.048	-0.102	-0.050	-0.066	-0.051	-0.019	-0.018	-0.030	-0.061	-0.030	-0.034	-0.042	-0.051	-0.039	-0.037	-0.049
	(0.104)	(0.048)	(0.043)	(0.039)	(0.014)	(0.011)	(0.008)	(0.006)	(0.026)	(0.002)	(0.002)	(0.002)	(0.003)	(0.003)	(0.003)	(0.003)
Female	-0.170	-0.115	-0.278	-0.183	-0.147	-0.147	-0.116	-0.126	-0.159	-0.215	-0.216	-0.231	-0.217	-0.197	-0.185	-0.216
	(0.076)	(0.047)	(0.049)	(0.050)	(0.027)	(0.027)	(0.019)	(0.015)	(0.050)	(0.006)	(0.005)	(0.006)	(0.007)	(0.006)	(0.006)	(0.006)
R^2	0.043	0.054	0.144	0.120	0.085	0.091	0.108	0.096	0.537	0.247	0.254	0.296	0.267	0.308	0.295	0.312
LAD																
S	0.027	0.030	0.041	0.038	0.033	0.034	0.038	0.033	0.048	0.056	0.065	0.068	0.065	0.075	0.079	0.083
	(0.011)	(0.009)	(0.006)	(0.006)	(0.005)	(0.004)	(0.003)	(0.003)	(0.007)	(0.001)	(0.001)	(0.001)	(0.002)	(0.001)	(0.001)	(0.001)
X	0.009	0.021	0.001	0.012	0.018	0.010	0.009	0.010	0.026	0.023	0.024	0.026	0.032	0.025	0.025	0.034
	(0.019)	(0.013)	(0.008)	(0.006)	(0.005)	(0.003)	(0.002)	(0.002)	(0.005)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
$X^2/100$	0.083	-0.100	-0.007	-0.026	-0.043	-0.022	-0.009	-0.015	-0.041	-0.034	-0.034	-0.039	-0.053	-0.037	-0.036	-0.054
	(0.102)	(0.052)	(0.028)	(0.020)	(0.015)	(0.009)	(0.006)	(0.005)	(0.015)	(0.003)	(0.002)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Female	-0.120	-0.091	-0.192	-0.058	-0.144	-0.131	-0.103	-0.091	-0.097	-0.194	-0.211	-0.224	-0.223	-0.193	-0.180	-0.206
	(0.066)	(0.053)	(0.034)	(0.032)	(0.027)	(0.021)	(0.015)	(0.015)	(0.035)	(0.006)	(0.006)	(0.007)	(0.007)	(0.006)	(0.006)	(0.007)
Pseudo R^2	0.021	0.047	0.054	0.043	0.057	0.085	0.072	0.063	0.435	0.154	0.154	0.173	0.156	0.179	0.173	0.185
N	459	609	854	805	1237	1339	1676	2606	1228	25565	23644	23919	15508	12518	18963	17486

Note: Dependent variable is ln(net monthly wage). N is the number of observations. For the period 1950–1993, a quadratic monthly time trend is included, and for the period 1994–2000, monthly dummies are included, to control for wage inflation. Standard errors are shown in parentheses.

Table 4: Variation in the Return to Schooling, by Region

	1994	1995	1996	1997	1998	1999	2000
S	0.065	0.074	0.080	0.076	0.082	0.084	0.092
	(0.002)	(0.003)	(0.003)	(0.004)	(0.003)	(0.003)	(0.004)
*S*North-East*	0.001	-0.003	-0.011	0.002	-0.003	-0.004	-0.010
	(0.004)	(0.004)	(0.004)	(0.006)	(0.004)	(0.005)	(0.005)
*S*South-East*	-0.013	-0.014	-0.014	-0.010	-0.009	-0.006	-0.009
	(0.003)	(0.004)	(0.004)	(0.005)	(0.004)	(0.005)	(0.005)
*S*South*	-0.013	-0.015	-0.014	-0.014	-0.004	-0.006	-0.004
	(0.003)	(0.004)	(0.004)	(0.005)	(0.004)	(0.005)	(0.005)
*S*South-West*	-0.008	-0.006	-0.023	-0.011	0.000	-0.003	0.004
	(0.004)	(0.005)	(0.004)	(0.006)	(0.005)	(0.005)	(0.006)
*S*West*	-0.011	-0.012	-0.020	-0.016	-0.005	-0.006	-0.015
	(0.004)	(0.005)	(0.005)	(0.006)	(0.005)	(0.005)	(0.006)
*S*North-West*	-0.003	-0.003	-0.015	-0.007	-0.004	0.007	-0.013
	(0.003)	(0.004)	(0.004)	(0.005)	(0.004)	(0.005)	(0.005)
*S*Center*	-0.007	-0.011	-0.017	-0.018	-0.013	-0.010	-0.020
	(0.003)	(0.004)	(0.004)	(0.005)	(0.004)	(0.005)	(0.005)
North-East	-0.112	-0.097	-0.001	-0.182	-0.105	-0.106	0.008
	(0.043)	(0.052)	(0.051)	(0.070)	(0.051)	(0.060)	(0.064)
South-East	0.165	0.130	0.131	0.084	0.041	-0.028	0.055
	(0.039)	(0.051)	(0.047)	(0.061)	(0.051)	(0.058)	(0.066)
South	0.098	0.085	0.079	0.090	-0.024	-0.058	-0.063
	(0.040)	(0.049)	(0.047)	(0.063)	(0.052)	(0.056)	(0.065)
South-West	0.061	0.017	0.211	0.064	-0.063	-0.076	-0.104
	(0.049)	(0.056)	(0.054)	(0.073)	(0.058)	(0.065)	(0.071)
West	0.135	0.147	0.203	0.123	-0.038	-0.031	0.122
	(0.052)	(0.056)	(0.062)	(0.072)	(0.061)	(0.062)	(0.073)
North-West	0.031	-0.027	0.111	0.034	-0.042	-0.210	0.062
	(0.042)	(0.050)	(0.049)	(0.062)	(0.051)	(0.057)	(0.063)
Center	-0.013	0.049	0.097	0.123	0.049	-0.029	0.131
	(0.041)	(0.051)	(0.048)	(0.062)	(0.050)	(0.056)	(0.066)
R^2	0.256	0.262	0.304	0.276	0.314	0.305	0.318

Note: Standard errors are shown in parentheses. Regions are defined on the basis of National Commission for Statistics (2000, p. 601). The equations also contain the other variables shown in Table 3 and monthly dummies to control for general wage inflation. Other variables are defined in Tables 1 and 2.

Table 5: Variation in the Return to Schooling, by Ethnicity

	1994	1995	1996	1997	1998	1999	2000
S	0.059	0.067	0.068	0.070	0.079	0.083	0.087
	(0.001)	(0.001)	(0.001)	(0.002)	(0.001)	(0.001)	(0.002)
*S*Hungarian*	-0.002	-0.001	-0.008	-0.016	-0.009	-0.004	-0.020
	(0.004)	(0.004)	(0.004)	(0.006)	(0.004)	(0.005)	(0.005)
*S*German*	-0.007	-0.012	0.008	0.007	-0.043	0.002	0.002
	(0.012)	(0.012)	(0.013)	(0.020)	(0.022)	(0.017)	(0.026)
*S*Roma*	-0.024	-0.030	-0.022	-0.013	-0.037	-0.052	-0.039
	(0.013)	(0.010)	(0.010)	(0.011)	(0.013)	(0.011)	(0.018)
*S*Other*	-0.037	-0.018	-0.029	-0.015	-0.035	0.000	-0.024
	(0.014)	(0.012)	(0.011)	(0.017)	(0.012)	(0.014)	(0.013)
Hungarian	-0.027	-0.058	0.011	0.116	0.066	-0.009	0.172
	(0.045)	(0.044)	(0.048)	(0.070)	(0.050)	(0.063)	(0.063)
German	0.065	0.102	-0.107	-0.118	0.514	-0.083	0.162
	(0.141)	(0.154)	(0.164)	(0.255)	(0.265)	(0.200)	(0.305)
Roma	0.018	0.125	0.140	0.092	0.224	0.475	0.336
	(0.089)	(0.086)	(0.076)	(0.095)	(0.118)	(0.097)	(0.163)
Other	0.420	0.239	0.331	0.202	0.398	-0.056	0.266
	(0.135)	(0.126)	(0.131)	(0.216)	(0.156)	(0.162)	(0.153)
R^2	0.249	0.256	0.298	0.268	0.309	0.297	0.314

Note: Standard errors are shown in parentheses. The equations also contain the other variables shown in Table 3 and monthly dummies to control for general wage inflation. Variables are defined in Tables 1 and 2.

Table 6: Variation in the Return to Schooling, by New versus Old Education

	1994	1995	1996	1997	1998	1999	2000
S	0.059	0.067	0.068	0.069	0.078	0.082	0.086
	(0.001)	(0.001)	(0.001)	(0.002)	(0.001)	(0.001)	(0.002)
*S*NEW*	0.021	0.022	0.010	0.003	0.007	0.004	0.000
	(0.009)	(0.007)	(0.006)	(0.006)	(0.004)	(0.004)	(0.004)
NEW	-0.352	-0.357	-0.215	-0.100	-0.172	-0.127	-0.083
	(0.114)	(0.085)	(0.073)	(0.074)	(0.054)	(0.049)	(0.051)
R^2	0.248	0.255	0.297	0.267	0.308	0.295	0.312

Note: Standard errors are shown in parentheses. The equations also contain the other variables shown in Table 3 and monthly dummies to control for general wage inflation. Variables are defined in Tables 1 and 2.

Table 7: **Variation in the Return to Schooling, by Sector**

	1994	1995	1996	1997	1998	1999	2000
S	0.055	0.062	0.066	0.067	0.076	0.080	0.084
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
*S*AGRIC*	0.005	-0.003	-0.009	0.000	0.002	0.003	-0.001
	(0.003)	(0.003)	(0.004)	(0.005)	(0.004)	(0.004)	(0.006)
*S*SERVICES*	0.011	0.012	0.010	0.010	0.008	0.005	0.008
	(0.002)	(0.002)	(0.002)	(0.003)	(0.002)	(0.003)	(0.003)
AGRIC	-0.229	-0.194	-0.159	-0.269	-0.273	-0.240	-0.212
	(0.037)	(0.036)	(0.043)	(0.056)	(0.049)	(0.054)	(0.069)
SERVICES	-0.208	-0.237	-0.275	-0.266	-0.229	-0.123	-0.190
	(0.024)	(0.025)	(0.027)	(0.035)	(0.029)	(0.031)	(0.036)
R^2	0.261	0.275	0.327	0.296	0.332	0.306	0.326

Note: Standard errors are shown in parentheses. The equations also contain the other variables shown in Table 3 and monthly dummies to control for general wage inflation. Variables are defined in Tables 1 and 2.

Table 8: Variation in the Return to Schooling, by Ownership of Employer

	1994	1995	1996	1997	1998	1999	2000
S	0.057	0.063	0.063	0.063	0.072	0.075	0.077
	(0.001)	(0.001)	(0.001)	(0.002)	(0.001)	(0.002)	(0.002)
*S*PRIVATE*	0.018	0.017	0.022	0.022	0.015	0.017	0.014
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
*S*MIXED*	0.004	0.007	-0.003	0.005	0.005	0.003	-0.005
	(0.007)	(0.006)	(0.005)	(0.006)	(0.004)	(0.004)	(0.005)
*S*COOP*	0.028	0.011	0.016	0.048	0.004	0.001	-0.001
	(0.009)	(0.010)	(0.013)	(0.016)	(0.013)	(0.013)	(0.015)
*S*OTHER*	0.046	0.013	0.011	-0.008	-0.049	-0.001	0.010
	(0.018)	(0.012)	(0.018)	(0.023)	(0.019)	(0.005)	(0.005)
PRIVATE	-0.257	-0.251	-0.339	-0.338	-0.282	-0.271	-0.269
	(0.038)	(0.034)	(0.035)	(0.042)	(0.032)	(0.032)	(0.036)
MIXED	-0.036	-0.056	0.065	-0.026	-0.029	-0.023	0.088
	(0.075)	(0.063)	(0.057)	(0.076)	(0.049)	(0.048)	(0.059)
COOP	-0.638	-0.445	-0.520	-0.839	-0.320	-0.216	-0.283
	(0.099)	(0.116)	(0.147)	(0.190)	(0.146)	(0.149)	(0.179)
OTHER	-0.733	-0.361	-0.350	-0.226	0.295	-0.029	-0.169
	(0.223)	(0.153)	(0.215)	(0.264)	(0.245)	(0.071)	(0.069)
R^2	0.258	0.264	0.308	0.278	0.322	0.303	0.325

Note: Standard errors are shown in parentheses. The equations also contain the other variables shown in Table 3 and monthly dummies to control for general wage inflation. Variables are defined in Tables 1 and 2.

Table 9: Return to Schooling for the Nonagricultural Self-Employed

	1994	1995	1996	1997	1998	1999	2000
S	0.072	0.095	0.137	0.132	0.155	0.137	0.127
	(0.013)	(0.013)	(0.009)	(0.011)	(0.009)	(0.008)	(0.008)
R^2	0.158	0.181	0.223	0.237	0.265	0.234	0.239
N	801	789	1310	861	1441	1492	1548

Note: These are coefficients on *S* (with standard errors in parentheses) from estimation of equation (1) by year for the nonagricultural self-employed respondents in the IHS.

Determinants of Household and Labor Income in the Baltic States:
Soviet and Post-Soviet Results

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Abstract

World Values Survey data are used to examine household income in the Baltic republics (Estonia, Latvia, and Lithuania) of the former Soviet Union. The household level data, gathered Summer 1990 (approximately one year prior to independence) by republic, provide a rare opportunity to empirically examine microeconomic factors influencing a Soviet household's position in the regional income distribution.

The Soviet-era results on income determination are compared with results on labor earnings using contemporary (1997 – 1999) labor force survey data gathered in each of the individual Baltic States. Particular attention is paid to how returns to human capital have changed during the transition from planned to market economy and on the changes in the distribution of income and wages across occupational and ethnic groups between the Soviet and post-Soviet periods. Specifically, the results indicate considerable increases in returns to education, a significant increase in returns to age/experience, a substantial increase in occupational wage dispersion, and a large shift in ethnic income differentials. The 1990 results indicate, accounting for a host of human capital, regional, occupational, and other factors, that ethnic Russians generally had equal or significantly higher (equal in Estonia – higher in Latvia and Lithuania) household incomes than did native Baltic residents. However, the contemporary labor force survey data indicate considerable unexplained ethnic wage differentials favoring native Baltic citizens in Estonia and Latvia (again accounting for a host of factors) and a roughly equal ethnic distribution in Lithuania. It is interesting to note that the assimilation of large Russian minorities in Estonia and Latvia has caused considerable political turmoil. Conversely, Lithuania has assimilated its relatively small Russian minority without strife.

JEL Classification: D31, P23, J71
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Introduction

A considerable literature on income in the Soviet Union exists (see Bergson (1984) for an excellent survey). However, due to a scarcity of data, microeconometric studies of income determination in the Soviet Union are rare. There is also a large and growing literature on income in the transition economies of the former Soviet Union (fSU) and Central and Eastern Europe (CEE) including a wealth of microeconometric studies. Much of this work has focused on CEE countries, and work on the fSU has tended to focus on Russia. The Baltic States of Estonia, Latvia, and Lithuania, granting some exception for Estonia, have been largely neglected in the literature. Again, this probably has much to do with data availability. This paper addresses these two gaps in the literature by presenting empirical evidence on determinants of household income distribution in the Baltic States while they were still republics of the Soviet Union (Summer 1990) and as independent states in the transitional period of the later 1990s.

Existing work on income in the Soviet Union has concentrated on income distribution and on savings – particularly forced savings. Bergson notes that studies of income inequality in the Soviet Union have been hampered by a lack of data. Despite this, he concludes that evidence indicates the level of inequality in the late 1970s into the early 1980s was considerably less than that of the U.S., but comparable to that of the Scandinavian countries. Bergson also notes that considerable fluctuations in income distribiution occurred in the Soviet coinciding with economic reforms enacted to stimulate productivity and growth (such as occurred when Khrushchev replaced Stalin).

The literature on savings tends to address motivations for saving in the Soviet Union testing the notion of whether or not there was monetary overhang (forced saving) in the Soviet Union in response to goods shortages. Earlier studies (Pickersgill (1976) and Ofer and Pickersgill (1980)) conclude that Soviet saving functions were actually quite similar to those of Western countries thus discounting the forced saving notion. However, using data released recently, Kim (1997 and 1999) provides evidence indicating monetary overhang was responsible for much Soviet household saving – particularly in the mid-late 1980s. While these studies all utilized measures of income and expenditures, none actually explored income determination in the Soviet Union.

Work on income in the transition economies of the fSU and CEE is extensive and fairly broad in scope. Atkinson and Mickelwright (1992) and Milanovic (1998) present excellent overviews of issues in income distribution, inequality, and poverty. Not surprisingly, both works find that levels of inequality and poverty have tended to increase sharply in the early years of transition. Much of this increase in equality relates to increased dispersion of labor earnings as economies in the fSU and CEE move away from the wage grids in place under central planning to wage distributions determined by market forces.

Several studies have examined income and labor earnings under socialist wage grids (see Smith (2001) for the Soviet Union and Munich et al. (2000) and Flanangan (1998) for Czechoslovakia). These studies find that wage distributions within the grids had narrow dispersions across several key categories related to education, experience, and occupation. Most studies find that wage dispersion across educational groups and occupational categories increases dramatically in the early years of transition (see Brainerd (1998), Newell and Reilly (1999), Orazem and Vodopivec (1995), and Rutkowski (1996) for examples). Evidence on returns to experience during transition is more muddled. Some have found decreased returns to work experience while others provide evidence of some increase in returns to experience. Additionally, evidence indicates that while absolute gender wage dispersion has increased, along with general wage dispersions, relative gender wage differentials have changed remarkably little between the planned and transitional periods in most of the fSU and CEE.

As mentioned, the Baltic States have been largely neglected in empirical work on wages and income. Exceptions are Kroncke and Smith (1999) who focused on ethnic wage differentials in Estonia and Noorkoiv et al. (1998) who conducted a general empirical study of wage and employment dynamics in Estonia from 1989 through 1995. Both papers use the retrospective nature of the first Estonian Labor Force Survey (covering the years 1989 – 1995) to examine changes in Estonian labor earnings in the late Soviet – early transitional period. Both studies indicate Estonian wages adjusted to market conditions in a manner similar to other transitional economies. Additionally, Kroncke and Smith present evidence indicating a lack of unexplained wage differentials across the primary ethnic groups in Estonia – ethnic Estonians and ethnic Russians – in 1989. However, a substantial unexplained wage differential favoring ethnic Estonians existed in 1994 relative to the now minority ethnic Russian group.

All three Baltic States are included in this study. Using micro-level data, determinants of household income in the late Soviet period and labor income in the transition period are examined from a comparative perspective. Particular is paid to the adjustment process from wage grid to market with respect to effect on income and wage differentials across groups with different levels of human capital, across occupational groups, and between the genders. Given the turmoil of assimilating large Russia minorities into Estonian and Latvian societies as opposed to the relative ease of assimilating the small Russian minority into Lithuanian society, ethnic differentials are also examined.

Overview of Data and Methodology

Microeconomic survey data from the late Soviet period (Summer 1990) and the late 1990s (1997-1999) are used to examine determinants of household income in the late Soviet period and labor earnings in the transitional period of each of the Baltic States. The 1990 data are from the World Values Survey (WVS). The WVS is an extensive survey conducted in 43 different regions and nations including the three Baltic States. The survey covers a broad range of topics related to politics, family life, religion, etc. It also contains rare Soviet-era data on income, occupation, education, and other variables that might be used to empirically examine income determination under the still heavily centralized Soviet system. Labor force surveys from each of the Baltic States are used to examine determinants of labor earnings in the later transition period.

While providing a rare opportunity to empirically examine income determinants in the Soviet Union, the WVS samples are relatively small. Further, the surveys were developed and conducted by noneconomists. Thus the data are not specifically structured for use by economists. The income variable is “total household income” in Lithuania and “total per capita household income” in Estonia and Latvia. To obtain the most relevant results, the samples for the World Values Survey are restricted to individuals who were the primary income earner for their household and were employed. People place themselves in income categories rather than give a specific ruble income. Thus the data are suitable for examining determinants of placement in the overall income distribution as opposed to estimating a standard Mincerian wage/income equation, and ordered logit equations are estimated to determine how factors commonly used in wage equations influenced standing in the Soviet income distribution.

The statistical offices of all three Baltic States began conducting labor force surveys in the mid 1990s. These surveys are generally similar to Western surveys. Thus they are fairly rich in detail and generally include relatively large samples. However, only the Estonian survey contains specific labor earnings data. Like the WVS, the Latvian and Lithuanian surveys only include wage categories. Therefore, as with the World Values Survey data, ordered logits are estimated to examine how various factors influence a worker's standing in the labor income distribution. To make the results as comparable as possible, an ordered logit is estimated for the Estonian sample as well. While differences in survey structure make comparisons between the Soviet-era WVS results and the contemporary Labor Force Survey results difficult, using ordered logit estimation for both does at least make for the most efficacious comparisons.

Summary of Preliminary Results

The results – particularly those using the labor force surveys – are still preliminary. The labor force surveys allow for more detailed estimations than have been conducted thus far. Additionally, estimations thus far are for a single period only. In each case data from two or three labor force surveys covering at least a two-year period are available for each Baltic State. Particular attention is paid to how certain economic characteristics, including human capital factors (education and experience – proxied by age in Estonia and Latvia) and occupation, and how certain personal characteristics, gender and ethnic group, affect income distribution.

Tables 1-4 present descriptive statistics of household (Tables 1 and 3) and labor (Tables 2 and 4) income distribution. Tables 1 and 2 present income distribution by ethnic group – native Baltic ethnicity and Russian ethnicity. In Estonia and Latvia, the ethnic distribution has clearly undergone considerable change. In 1990, Estonia had a fairly even ethnic income distribution. However ethnic Estonians clearly fare better in the 1997 distribution. The 1990 income distribution in Latvia clearly favored ethnic Russians while the 1998 labor income distribution clearly favors ethnic Latvians. In comparison any change between the 1990 and 1999 Lithuanian distributions seems minor. The difference in the evolution of the income distribution by ethnic group is potentially interesting in light of the situation in the three Baltic States. As indicated by Tables 1 and 2, Estonia and Latvia have relatively large ethnic minorities. Both have faced considerable political and social turmoil related to the assimilation of their ethnic Russian minorities. Conversely, Lithuania, with a small Russian minority, has faced relatively little trouble assimilating its Russian minority. Thus considerable attention is paid to ethnic differentials.

Conversely, the gender distributions show relatively little change over time. In all three Baltic States households headed by females fared relatively poorly at the end of the Soviet period, and women fare relatively poorly in the labor income distributions of all three in the later 1990s.

Tables 5 and 6 present the preliminary ordered logit results for 1990 and the later 1990s respectively. While it is recognized that comparisons of results across time or across countries should be viewed with caution, some interesting differences are apparent and will be considered in detail in the future. In sum, it is clear that age/experience and education have clearly become more important during the transition period in the Baltic States. Further, occupational differences have become considerably more pronounced. This is consistent with findings from other transitional economies.

With respect to the effect of age on earnings, the 1990 results are somewhat contradictory (see Table 5). They range from a significant negative effect in Estonia to an insignificant effect in Latvia, to a significant positive effect in Lithuania. The Lithuanian results may have something to do with the absence of data on location (not available for Lithuania in the 1990 data). The results from the transition era indicate age effects are much more potent in the capital cities than outside the capitals (see tables 6 and 7). In the late 1990s (Table 6) the results for Estonia and

Latvia both show evidence of significant age effects that would tend to favor older workers – in-line with what one would expect for a market economy. However, the results for Lithuania provide evidence that the most important aspect of experience in transition might be job-specific experience (tenure) as opposed to general work experience. Table 7 provides strong evidence that positive age effects are important in the capital cities of the Baltic States but are much less influential outside of the primary urban area. Though the reasons behind this remain to be explored further, a likely explanation lies in the importance of public sector jobs in the capital cities that are far more likely to use experience-based wage scales.

Consistent with findings elsewhere (see Bergson (1984) for example), the occupational wage distribution in the Soviet Baltic States is quite tight. In particular, blue-collar workers do relatively well. This result is also consistent with Soviet priorities regarding manufacturing relative to service work. Occupational wage differentials have increased in the transitional Baltic States and have tended to move towards a distribution more typical of a market economy. Specifically, skilled white-collar workers are now clearly at the top of the wage ladder and unskilled blue-collar work is quite poorly paid. The later results indicate a severely depressed agricultural sector though Latvia is somewhat of an exception in this regard.

While gender results, in a qualitative sense, have changed little, controlling for a variety of other variables, the results do indeed indicate a substantial change in the effect of ethnicity on standing in the income distribution in Estonia and Latvia (see Table 6). In Estonia, the results indicate no statistical effect of ethnicity on household income in 1990. However, there is a strong and significant effect in the 1997 estimation. In Latvia, there is a complete reversal of the effect of ethnicity on income distribution. In 1990, controlling for all variables present in Table 5, ethnic Russian households fared considerably better than ethnic Latvian households. However, in 1998, ethnic Latvians, all else equal, earn considerably more labor income than do ethnic Russians. In Lithuania, there is little evidence of an ethnic household income or labor income effect in either sample (at least when examining the entire Lithuanian sample).

Further, there is a remarkable difference in the results with respect to ethnicity when the samples are separated by work within and outside the capital city. In all cases, including Lithuania, ethnic Russians fare relatively poorly within the capital city. However, outside the capital city, the full spectrum is covered with respect to wages and ethnic Russians. In Estonia there is no statistical difference between Estonians' standing in the wage distribution and ethnic Russians. Conversely, though the statistical difference appears to be smaller than within Riga, ethnic Russians are still at a significant disadvantage vis-à-vis ethnic Latvians outside of Riga. Finally in Lithuania, ethnic Russians actually tend to be higher in the wage distribution than their Lithuanian counterparts. This is another aspect of the results that requires further exploration. One possible answer lies again in the preponderance of public administration jobs in the capital cities that Russians may be shut out of due to citizenship and/or language requirements.

Concluding Remarks and Future Work

The preliminary results indicate the Baltic States, in many respects, have made rapid progress towards a wage/income distribution shaped by market forces. In particular, wage dispersions across occupational groups and across groups with different educational levels have widened dramatically. While the evidence on age/experience tends to be less clear in transition economies, the results presented here indicate that returns to experience may be increasing as well. However, the data from Lithuania indicate that more recent job-specific experience is considerably more influential than general work experience – much of which may have been gained during the Soviet period. Relative gender distributions appear to have changed little during the transition period. This is not surprising given relative gender differentials in the Soviet-era Baltic States were somewhat similar to those existing in most market economies. However, a considerable change has occurred with respect to ethnic wage/income differentials. In 1990, when Russians were the dominant ethnic group (at least in the Soviet Union as a whole), the results indicate that households headed by ethnic Russians did about as well as households headed by ethnic Balts in Estonia and Lithuania and considerably better than households headed by ethnic Latvians in Latvia. By the late 1990s, when Russians were a distinct ethnic minority, the situation had been reversed in Latvia, ethnic Estonians had gained a clear advantage in Estonia, but in Lithuania the situation remained essentially unchanged with no evidence of ethnic wage differentials. Interestingly enough, Lithuania is the only Baltic State to avoid significant ethnic tension in its transition period. The relatively small ethnic Russian minority (approximately ten percent of the population) has been assimilated into Lithuanian society with relative ease. Conversely, with quite large Russian minorities (just under 30 percent in Estonia and just over 30 percent in Latvia), Estonia and Latvia have faced considerable difficulty assimilating their ethnic Russian populations into society.

In future work more detail can be added to the estimations using recent labor force survey data. In particular, industry data are available for each Baltic labor force survey (though not for the WVS data). While changes in the Baltic labor force surveys (the statistical offices have little experience gathering survey data and continue to adjust the surveys over time) make it difficult to pool data, multiple surveys exist for each Baltic State. The results of these will be examined individually and, when possible, data will be pooled to increase the reliability of results.

Table 1
Income Distribution by Ethnicity - Summer 1990 (percentage in each group)

	Native Baltic Ethnicity			Russian Ethnicity		
Variable	Estonia	Latvia	Lithuania	Estonia	Latvia	Lithuania
N	366	235	324	191	196	49
inc1	15.85	27.66	9.26	18.85	18.88	6.12
inc2	25.96	29.36	10.19	22.51	23.98	14.29
inc3	26.78	17.02	26.54	26.70	21.94	20.40
inc4	19.95	20.43	34.57	21.99	19.39	34.69
inc5	11.48	5.53	19.44	9.95	15.82	24.49

Table 2
Income Distribution by Ethnicity (percentage in each group)

	Native Baltic Ethnicity			Russian Ethnicity		
Variable	Estonia (Jan. 97)	Latvia (May 98)	Lithuania (May 99)	Estonia	Latvia	Lithuania
N	1821	3439	3250	705	1448	246
inc1	19.66	9.01	21.14	20.28	8.01	7.32
inc2	18.34	21.69	13.88	22.84	22.17	13.01
inc3	19.82	38.99	16.28	24.68	40.75	15.85
inc4	19.06	26.14	13.82	18.30	24.38	22.76
inc5	23.12	4.16	13.02	13.90	4.70	15.04
inc6			10.74			13.82
inc7			11.14			12.20

Table 3
Income Distribution by Gender - Summer 1990 (percentage in each group)

	Male			Female		
Variable	Estonia	Latvia	Lithuania	Estonia	Latvia	Lithuania
N	350	284	277	251	238	167
inc1	14.29	20.77	5.05	19.52	28.99	17.37
inc2	23.14	25.35	9.03	27.89	26.89	16.12
inc3	25.43	17.25	24.55	26.98	18.49	28.74
inc4	22.29	21.83	35.74	18.73	16.81	29.34
inc5	14.86	14.79	25.63	6.77	8.82	8.38

Table 4
Income Distribution by Gender (percentage in each group)

	Men			Women		
Variable	Estonia	Latvia	Lithuania	Estonia	Latvia	Lithuania
N	1384	2771	1996	1346	2727	1880
inc1	16.18	7.76	20.19	23.55	9.90	18.78
inc2	16.04	18.12	10.37	24.44	25.96	17.77
inc3	19.80	36.74	13.73	21.92	42.46	18.94
inc4	21.03	31.25	14.33	17.24	19.44	15.05
inc5	26.95	6.13	14.28	12.85	2.24	12.23
inc6			13.03			9.52
inc7			14.08			7.71

Table 5
Ordered Logit Regression Results (Dependent varible household income)

Variable	Estonia (Summer 90)	Latvia (Summer 90)	Lithuania (Summer 90)	Variable Definition
Age	-0.091** (0.042)	-0.033 (0.044)	0.148*** (0.046)	Age in years
age^2	0.001*** (0.0005)	0.0007 (0.0005)	-0.002*** (0.0005)	Age squared
0=female 1=male	0.579*** (0.167)	0.370** (0.174)	1.206*** (0.204)	Gender dummy
Education	-0.005 (0.044)	0.085* (0.044)	-0.005 (0.045)	Years of education
Capital	1.200*** (0.239)	0.761*** (0.259)	--	Dummy for job location in the capital city
urban1	1.228*** (0.344)	0.466 (0.298)	--	Dummy for job location in cities w/pop. 100-500,000
urban2	0.800*** (0.254)	0.470 (0.319)	--	Dummy for job location in cities w/pop. 20-100,000
urban3	0.682** (0.278)	-0.098 (0.327)	--	Dummy for job location in cities w/pop. 5-20,000
urban4	0.049 (0.317)	-0.420 (0.301)	--	Dummy for job location in cities w/pop. 2-5000
Hours (full-time=1)	0.619* (0.331)	-0.858*** (0.273)	0.486*** (0.184)	Dummy for full-time work (over 30 hours per week)
native Balt	0.205 (0.179)	-0.531*** (0.184)	-0.107 (0.297)	Dummy for native Baltic ethnicity
Pole	--	--	0.941** (0.465)	Dummy for Polish ethnicity
Other	0.200 (0.309)	-0.131 (0.238)	-1.208*** (0.441)	Dummy for ethnicity other than native Baltic, Russian (the reference group) or Polish (Lithuania only)
Self-employed	1.633** (0.692)	0.456 (0.502)	1.524*** (0.590)	Occupational dummy for self-employment (unskilled labor is the reference occupation)
Manager	1.061*** (0.368)	0.779* (0.430)	1.617*** (0.515)	Management occupation dummy
Professional	0.413 (0.346)	0.886** (0.427)	0.739** (0.351)	Dummy for a professional occupation
Wcow	0.508* (0.293)	0.376 (0.415)	0.251 (0.241)	Dummy for white collar office workers
Skbc	0.222 (0.272)	0.851** (0.414)	0.451 (0.313)	Dummy for skilled blue collar occupations
Ag	0.818 (1.382)	-0.098 (0.665)	-0.507 (0.449)	Dummy for agricultural workers
log likelihood	-902.29	-764.69	-606.61	
$chi^2(k)$	82.44***	94.82***	92.78***	

Table 6
Ordered Logit Results (Dependent Variable: labor income)

Variable	Estonia (Jan. 97)	Latvia (May 98)	Lithuania (May 99)	Lithuania (w/o urban centers)	Variable Definitions
Age	0.087*** (0.019)	0.070*** (0.014)	--	--	Age in years
age^2	-0.001*** (0.0002)	-0.0009*** (0.0002)	--	--	Age squared
Tenure	--	--	0.064*** (0.011)	0.056*** (0.010)	Years of experience in current job
$Tenure^2$	--	--	-0.0008** (0.0003)	-0.0007** (0.0003)	Tenure squared
Exp	--	--	0.009 (0.008)	0.007 (0.008)	Years of work experience outside of current job
exp^2	--	--	0.0004 (0.0002)	0.0005** (0.0002)	Experience squared
0=female 1=male	0.804*** (0.082)	0.937*** (0.059)	0.785*** (0.068)	0.766*** (0.068)	Gender dummy
Education	0.211*** (0.029)	0.186*** (0.019)	0.249*** (0.021)	0.264*** (0.021)	Educational level (primary, secondary, ..., higher)
Capital	0.787*** (.084)	1.533*** (0.074)	0.708*** (0.090)	--	Dummy for job location in the capital city
urban1	0.236* (0.126)	0.868*** (0.064)	0.304*** (0.092)	--	Dummy for job location in the 2nd largest city
urban2	-0.179 (0.174)	--	0.793*** (0.128)	--	Dummy for job location in the 3rd largest city
urban3	0.201 (0.197)	--	0.621*** (0.146)	--	Dummy for job location in the 4th largest city
urban4	--	--	-0.256* (0.156)	--	Dummy for job location in the 5th largest city
hours (ft=1 est)	1.146*** (0.074)	0.024*** (0.002)	0.029*** (0.004)	0.030*** (0.004)	Hours worked per week
native Balt	0.405*** (0.090)	0.290*** (0.061)	-0.037 (0.125)	-0.298** (0.120)	Dummy for native Baltic ethnicity
Belarus	--	-0.036 (0.140)	-0.308 (0.267)	-0.272 (0.264)	Dummy for Belarussian ethnicity
Ukrainian	--	-0.253 (0.172)	1.242*** (0.439)	1.104** (0.436)	Dummy for Ukrainian ethnicity
Pole	--	--	-0.272* (0.162)	-0.224 (0.159)	Dummy for Polish ethnicity
Other	-0.065 (0.148)	0.232 (0.224)	-0.502 (0.390)	-0.615 (0.388)	Dummy for other ethnic groups (Russians are the reference group)
Manager	1.916*** (0.174)	2.423*** (0.127)	2.394*** (0.146)	2.424*** (0.146)	Management occupation dummy
Professio nal	1.860*** (0.177)	1.679*** (0.108)	2.255*** (0.144)	2.263*** (0.143)	Dummy for a professional occupation
Technical	1.285*** (0.158)	1.305*** (0.100)	1.426*** (0.151)	1.432*** (0.150)	Dummy for workers in techinical occupations

Table 6 (cont.)

Clerical	1.117*** (0.201)	1.215*** (0.123)	1.172*** (0.158)	1.152*** (0.158)	Dummy white collar clerical workers
Service	0.213 (0.164)	0.404*** (0.096)	0.643*** (0.127)	0.681*** (0.127)	Dummy workers in service occupations
Craft	1.114*** (0.150)	1.008*** (0.092)	0.903*** (0.113)	0.908*** (0.112)	Dummy for skilled craft workers
Skbc	0.716*** (0.160)	0.920*** (0.097)	0.969*** (0.133)	0.969*** (0.133)	Dummy for skilled blue collar workers
Ag	-0.457** (0.223)	0.162 (0.185)	-2.207*** (0.136)	-2.394*** (0.133)	Dummy for agricultural workers (unskilled workers are the reference group)
log likelihood	-3841.90	-6751.21	-5986.02	-6036.89	
$chi^2(k)$	1101.70***	1841.51***	2975.24***	2873.50***	

Table 7
Ordered Logit Results (Dependent Variable: labour income)

	Capital City			Outside Capital		
Variable	Estonia	Latvia	Lithuania	Estonia	Latvia	Lithuania
age	0.159*** (0.030)	0.109*** (0.024)	--	0.032 (0.025)	0.037** (0.017)	--
age^2	-0.002*** (0.0004)	-0.001*** (0.0003)	--	-0.0005 (0.0003)	-0.0004** (0.0002)	--
ten	--	--	0.038* (0.023)	--	--	0.043*** (0.011)
ten^2	--	--	-0.0002 (0.0007)	--	--	0.0001 (0.0004)
exp	--	--	0.048*** (0.017)	--	--	-0.0008 (0.008)
exp^2	--	--	-0.001 (0.0005)	--	--	0.0005 (0.0003)
0=female 1=male	0.993*** (0.129)	1.146*** (0.100)	1.020*** (0.135)	0.895*** (0.091)	0.891*** (0.062)	0.719*** (0.067)
education	0.296*** (0.038)	0.327*** (0.033)	0.460*** (0.039)	0.374*** (0.029)	0.354*** (0.020)	0.480*** (0.019)
urban1	--	--	--	0.390*** (0.128)	0.885*** (0.062)	0.862*** (0.090)
urban2	--	--	--	-0.494*** (0.183)	--	1.274*** (0.128)
urban3	--	--	--	0.464** (0.207)	--	1.232*** (0.146)
urban4	--	--	--	--	--	0.316** (0.151)
hours (ft=1 est)	1.925*** (0.223)	0.020*** (0.003)	0.036*** (0.008)	2.259*** (0.173)	0.024*** (0.002)	0.027*** (0.003)

Table 7 (cont.)

priv.=1 pub.=0	0.093 (0.134)	0.280*** (0.090)	-0.209 (0.141)	-0.370*** (0.094)	-0.345*** (0.058)	-0.930*** (0.075)
native Balt	1.250*** (0.138)	0.517*** (0.105)	0.548*** (0.186)	-0.053 (0.117)	0.315*** (0.075)	-0.329** (0.164)
Belarus	--	-0.033 (0.225)	-0.594* (0.341)	--	0.026 (0.176)	0.259 (0.402)
Ukrainian	--	-0.239 (0.256)	0.588 (0.679)	--	-0.169 (0.227)	1.424*** (0.543)
Pole	--	--	0.183 (0.205)	--	--	-0.806*** (0.266)
Other	0.041 (0.217)	-0.006 (0.202)	-0.415 (0.652)	-0.091 (0.201)	0.076 (0.151)	-0.524 (0.449)
log likelihood	-1243.0	-1931.7	-1419.1	-2546.0	-5039.3	-5959.4
$chi^2(k)$	333.99***	309.88***	289.88***	497.38***	800.75***	1636.28***

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Inter–regional Mobility in the Accession Countries: A Comparison to EU-Member States

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Abstract

This paper uses a data set covering 9 EU member states and 7 candidate countries to compare inter-regional migration patterns in the 1990s. We find that migration is lower in candidate countries than in EU member states. Also in contrast to the member states migration has fallen in candidate countries. This casts doubt on the viability of migration as an adjustment mechanism. Estimating place to place models of migration we also find that migration is less reactive to regional disparities in candidate countries than in EU member states. If reaction to labor market disparities were similar to EU states gross migration should increase by 10% to 50% and net migration by a factor by 2 to over 10.

Key Words: Regional Labor Market Adjustment, Transition, EU - Accession
JEL – Classification: P25, J61

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1 Introduction

The stylized fact of low migration rates in Europe has been repeatedly documented. Decressin/Fatas (1995), Fatas (2000), Obstfeld/Peri (2000) and Puhani (2001) find that migration only contributes moderately to the reduction of differences in regional labor market conditions in European Union (EU) – member states. Recent evidence suggests that migration is an even less efficient mechanism for equilibrating regional labor markets in candidate countries. Fidrmuc (2004) finds that overall internal mobility in candidate countries is low and inefficient in reducing regional disparities. Ederveen/Bardsley (2003) find that migrants in the candidate countries are less responsive to regional wage and employment disparities than in current EU member states and Drinkwater (2003) reports that the willingness to migrate across regions and national borders is at the lower end among European countries. Cseres-Gergeley (2002), Hazans (2003), Kallai (2003) and Fidrmuc/Huber (2003) provide case studies on Hungary, the Baltics, Romania and the Czech Republic to provide further evidence on low migration in candidate countries.

The potential economic and political consequences of this lack of labor mobility have been repeatedly stressed. Low internal migration increases mismatch unemployment and will thus contribute to high nation wide unemployment (Boeri/Scarpetta, 1996). Aside from causing social problems, this may also have political implications. In the long run higher unemployment rates may lead to increased demands for regional transfers. This in turn may cause dissatisfaction on the side of those parts of the population financing regional transfers and lead to the disintegration of political Unions.[2] Furthermore, lack of migration impinges on the short run adjustment capabilities of regional labor markets to asymmetric shocks (Eichengreen, 1998). Lacking migration may thus also hamper the viability of monetary Unions. Since exchange rate fluctuations are impossible in monetary Unions, the absence of migration, leads to adjustment to asymmetric shocks through wages, unemployment or participation rates. To the extent that these adjustment mechanisms are socially or politically less desirable than migration, low migration will cause social and political costs of EMU (Fidrmuc, 2003).

Despite these profound implications, little is known about the causes for low migration in Europe. A number of explanations such as inefficiencies in spatial matching (Faini et al, 1997), the effects of social transfers on the search incentives of the unemployed (Fredriksson, 1999), housing market imperfections (Cameron/Muellbauer, 1998) and cultural differences as reflected for instance in attitudes towards risk (Bentivogli/Pagano, 1999) have been put forward to account for this puzzle. A final verdict on which of these factors is decisive, however, has not been reached.

In this paper we use data on inter-regional migration in the 1990s for nine current EU – member states and seven countries that either joined the European Union in 2004 or are negotiating on membership, to compare regional migration patterns in candidate countries to those in the EU. Our goals are twofold. First, we explore the stylized facts of migration in candidate countries and compare them to EU member states. In the next section we thus describe migratory moves in the two regions. We highlight a number of differences in migration patterns. In particular interregional migration is low by EU standards in candidate countries and has been falling throughout the 1990s. A lower share of migration is accounted for by active aged persons and in both regions and around 90% of all measured migration flows are churning flows, which contribute little to the equilibration of aggregate regional disparities. We also present evidence that a substantial part of migration covers only short distances and that migration rates are strongly correlated over time. This suggests that migration presents a rather protracted and sluggish adjustment mechanism to regional disparities.

Second, we compare the responsiveness of migration to regional income and labor market disparities by estimating place to place models of migration. We estimate a model suggested by Bentivogli/Pagano (1999), incorporating risk aversion in section three. In contrast to earlier comparative work, this allows us to estimate directly the elasticity of migration with respect to regional income and employment rate disparities in both member states and candidate countries. We find that both net and gross migration is less reactive to regional employment rate and income disparities in the candidate countries and that attitudes towards risk play a minor, but geographic factors a major role in determining migration. We also show that gross migration should increase by 10% to 50% in candidate countries if it were as responsive to regional disparities in candidate countries as in Spain, Italy or the Netherlands. Net migration should increase by a factor of 2 to 10. Section four finally concludes the paper by drawing some policy conclusions and outlining potential directions for further research.

2 Stylized Facts

We use internal migration data for the 1990s on nine European Union countries namely, Austria, Belgium, Denmark, Germany, Italy, the Netherlands, Spain, Sweden and the UK and seven countries which either have completed negotiations for membership or are still negotiating on accession namely, the Czech Republic, Estonia, Hungary, Poland, Slovenia, Slovakia and Romania taken from Eorostat's Cronos database. As shown in table 1 these data vary in scope and content. In particular, the data refer to different regional units in various countries. For most countries data refer to NUTS II regions, but for Denmark, Estonia and Slovenia, data are available only at the NUTS III level, while in Germany and the UK they only cover NUTS I regions. These differences in regional disaggregation imply substantial differences in region size. For instance, the largest territories in terms of average population are the German and U.K. NUTS I regions and the smallest regions are the NUTS III regions of Slovenia, Estonia and Demark. For regional units at the same level of regional disaggregation average size also varies considerably. In terms of population the largest NUTS II regions are in Italy with 2.9 million inhabitants and the smallest in Austria with 898 thousand.

The data also differ with respect to the time period covered[3]. For Germany for instance data are only available to 1993 and in Slovakia only the year 2000 is available. Thus in an attempt to maximize available information, we conduct our descriptive analysis for two sample years: 1992 and 1999.[4] We break this rule only in the cases of Poland, where we report data from 1990 instead of 1992 and for Slovakia where data from the year 2000 are taken instead of 1999. Furthermore, most of the data collected are place to place data. For two countries (Romania and Slovakia), however, place to place information is not available.[5] Thus we cannot conduct analysis in the same depth for these countries.

{Table 1: Around here}

2.1 Net and Gross Migratory Moves

In Table 2 we report the number of migrants changing their region of residence as a percentage of the country's population in 1992 and 1999, respectively. This indicator has been used as a measure of the overall mobility by a number of authors (e.g. Fatas, 2000, Faini et al, 1997, Bentolila, 1999). Formally, it can be defined as half of the sum of total outflows and inflows across regions[6]:

$$GF = \frac{1}{2}\left[\frac{\sum_i (O_i + M_i)}{\sum_i POP_i}\right] \tag{1}$$

where GF stands for the share of gross migration flows in total population, O_i and M_i are the migrant outflows and inflows from region i, respectively, and POP_i is the population of region i.

Gross migration may, however, be a misleading indicator, because a substantial part of migration is accounted for by churning flows, where people move in and out of the same region.[7] Most macro-economic models, which consider migration as an equilibrating mechanism in the face of regional disparities focus on net-migration. Thus measures of net migration should better capture the efficiency of inter regional migration flows in equilibrating regional disparities in unemployment and income. This can be measured as the sum absolute values of the difference between emigration and immigration across regions. In the notation of equation (1) net migration flows as a share of total population are given by:

$$NF = \frac{1}{2}\left[\frac{\sum_i |O_i - M_i|}{\sum_i POP_i}\right] \tag{2}$$

Furthermore, from the above definitions of net and gross migration rates and noticing that:

$$\left[\frac{\sum_i (O_i + M_i)}{\sum_i POP_i}\right]\left[\frac{\sum_i |O_i - M_i|}{\sum_i (O_i + M_i)}\right] = \left[\frac{\sum_i |O_i - M_i|}{\sum_i POP_i}\right] \tag{3}$$

the share of net flows in total flows is:

$$SNF = \left[\frac{\sum_i |O_i - M_i|}{\sum_i (O_i + M_i)}\right] \tag{4}$$

The results of this decomposition (see table 2) do not suggest that migration is a viable mechanism for regional adjustment in Europe. Although there is some variance across countries, migration is low in current EU-member states and even lower in candidate countries. In the average EU member state around 1% of the population changes region of residence within a year. Gross migration rates are substantially lower than 1% only in Italy and Spain. In the candidate countries gross migration rates exceed the 1% mark only in Romania and Hungary and are around or below 0.5% in most countries.

{Table 2 around here}

Furthermore, in contrast to the EU-Member states, where gross migration has stagnated or even increased over the period from 1992 to 1999, migration rates have fallen in all candidate countries for which we have data in both time periods. This finding is consistent with a number of results reported by other authors researching migration patterns in the candidate countries (Kallai, 2004, Hazans, 2004, Fidrmuc/Huber, 2004) but stands in stark contrast to the increase of regional disparities found in much of the literature on regional development in the candidate countries (Egger/Huber/Pfaffermayr 2004, Petrakos 1995, Huber/Palme, 2001, Gorzelak, 1996), which suggests that regional divergence predominated in the last decade in the candidate countries and thus incentives to migrate should have increased rather than decreased.

The low effectiveness of migration at lowering regional disparities is underlined by net migration rates. They rarely exceed 0.1% of the population in the candidate countries and haven fallen in all countries but the Czech Republic.[8] In current EU-member states by contrast net migration flows at least approach the 0.1% level in all countries but Austria and the Netherlands and the evidence concerning a decline is less ubiquous. Thus a substantial part of migration (around 90%) in both regions is due to churning flows, which contribute little to the narrowing of aggregate regional disparities.

2.2 Regional and Demographic Structure

Our data refers to population moves. This may distort results concerning labor migration, if some migration is undertaken for reasons other than economic activity. Examples of such migration may be students moving to their place of education or pensioners to retire. Furthermore, as noted for example by Cameron/Muellbauer (1998) migration among neighboring regions and within urban agglomerations may be primarily motivated by housing motives, if residents of one region (such as a city) move to another (such as the suburbs) without changing workplace. Such migration is obviously not associated with income or unemployment disparities between regions, but is motivated by cheaper housing, better educational infrastructure or better living conditions in the receiving region. Thus it will do little to equilibrate regional labor market disparities, since effective labor supply remains unchanged both in the sending and receiving region.

{Table 3 Around here}

While gauging the exact extent of such non-labor market motivated migration is impossible with our data, some indication is available. First, for a number of countries we have available migration by age groups and gender.[9] This allows us to estimate the share of active aged (between 20 and 64) in total migration i.e., of those that at least theoretically could move for labor market reasons. These data (see Table 3) suggest that the share of active aged is slightly lower in most candidate countries than in the EU member states. In typical candidate countries between 65% and 70% of the migrants are active aged, (with the outliers being Romania with 74% and Estonia with around 58%). In the member states by contrast typically more than 70% of the migrants are active aged. The only indicator, where candidate countries have higher figures than member states is with the share of female migrants. More than half of the migrants in candidate countries are female. This may in part be explained by the higher participation rate of females in many candidate countries, leading to more labor motivated migration among women.

{Table 4 Around Here}

Furthermore, for those countries where place to place data are available we can calculate the share of moves between neighboring regions as indication of the relevance of short distance moves, which are not associated with labor market motives. Shares of migration among neighboring regions may, however, be influenced by differences in geography among countries, which may in turn lead to differences in the number of neighbor relationships and thus may influence the share of migration between neighboring regions. In column 3 of table 3 we thus calculated the share of contingency relationships in a country.[10] Comparing this share with the share of migration among neighboring regions gives an indication of the extent to which the share of short distance moves between neighboring regions exceeds the rate expected if migration were independent of distance. According to these statistics flows between neighboring regions exceed their expected value by a factor of between 1.2 and 3.0. Thus a substantial part of migration in both candidate countries and EU member states is accounted for by short distance moves.[11]

Further doubt concerning the viability of migration as a mechanism for equilibrating regional disparities comes from correlating net migration rates (as a percentage of resident population in a region) over two time periods. These correlations are usually high and significant (see column 4 of table 4). Correlation coefficients of net migration rates between regions at two points in time seven years apart are highly significant in all countries and may reach levels of up to 0.9. As recently pointed out by Rappaport (1999) this suggests that migration is not reactive to transitory shocks but reflects either the protracted adjustment to permanent shocks or differences in the steady state growth paths among regions.

2.3 Internal and External Migration

Our data also exclusively measure internal migration. A number of recent contributions, however, suggest that international and intra-country migration may be substitutes (Borjas, 1999). If migrants from abroad are more likely to move to places with high wages and low unemployment rates, this may deter national migrants from moving to these places. Alternatively if emigrants in depressed regions are faced with a choice of moving to less depressed regions in their own country or abroad, the choice may be to move abroad, if these regions offer even better conditions than regions at home.

{Table 5 Around here}

Again this claim can be analyzed at least for a subset of countries in our data, for which we have available information on net migration abroad from the same data set. The information displayed in table 5, suggests a low potential for this explanation. While most candidate countries (except for Estonia) are net receiving countries for international migrants the share of migrants received tends to be low.

Similarly, emigration abroad does not seem to be a viable alternative to migration within a country. Most of the candidate countries for which data are available, have gross emigration rates abroad that are at the lower end of the EU distribution.[12]

Finally regional data suggests that rather than substitutes international migration is complementary to internal migration. Regions with high net emigration into the country also tend to be regions with high emigration abroad. The correlation coefficient between the two is 0.45. Thus, it seems unlikely that high international migration rates compensate for low internal migration in candidate countries

3 Estimating Place to Place Models of Migration

Descriptive statistics thus suggest that migration rates in the candidate countries are low even relative to EU figures and have fallen in the last decade. Furthermore, they indicate that a larger share of migration is accounted for by population moves not associated with labor market motives and that migration is highly auto-correlated. While this indicates that migration may be ineffective in reducing labor market disparities, it does not provide us with quantitative estimates. We therefore estimate a model of place-to-place migration to quantify differences in the responsiveness of migration to regional disparities. To motivate our choice of specification, we consider a model proposed by Bentivogli/Pagano (1999). In this overlapping generations model, agents are assumed to live for two periods. At the beginning of the first period they decide, whether they would like to live in their region of birth (labeled h) or whether they prefer emigration to another region (called a) within the country. After this decision has been made agents in their first period consume in their chosen region of residence and either work receiving income of w_t, which is drawn from a normal distribution with mean μ_i and variance σ_i (with i an index for the region of residence i.e. $i \in \{a, h\}$,) or are registered as non-employed and receive an income from the informal sector of b, which is assumed constant across all regions. Finally, in their second period of life agents retire and consume from their savings.

If agents at the beginning of the first period decide to emigrate from their region of birth they incur a cost of migration, denoted by θ_{ah}. Bentivogli/Pagano (1999) show that under the assumption that θ_{ah} is uniformly distributed in the interval [p,z] (where p depends on the relative attractivity of regions as well as the costs of migration) among agents, the share of population of a region moving from region a to h at time t (m_{aht}) can be written as:

$$m_{aht} = \alpha \ln(\mu_{at} - \mu_{ht}) - \alpha b \ln(u_{at} + u_{ht}) + \frac{\alpha\lambda}{2} \ln(\sigma_a{}^2 - \sigma_h^2) - p_{ah} \tag{5}$$

with α a function of the interest rate, and λ the absolute risk aversion coefficient and u_{it} and σ_{it} indicators of labor market tightness and the variance of regional income, respectively.

In empirically implementing equation (5) we include fixed effects to control for time invariant characteristics of regions such as amenities as well as psychological and financial costs associated with migration. In particular we reformulate equation (5) as:

$$m_{aht} = \alpha \ln(\mu_{at} - \mu_{ht}) - \beta \ln(u_{at} + u_{ht}) + \gamma \ln(\sigma_{at}{}^2 - \sigma_{ht}^2) - \sum_a \sum_{h \neq a} \phi_{ah} + \sum_t \tau_t + \varsigma_{aht} \tag{6}$$

where ϕ_{ah} is a set of *Jx(J-1)* fixed effects for each sending and receiving region pair. These are included to control for all aspects of moving costs between two regions, e.g., the differences in regional amenities, the distance to be covered, contingency effects, differences in relationships between urban and suburban regions, and potential cultural differences within regions of countries that may increase psychological moving costs. τ_t are fixed effects for each time, period. These are included to proxy for macroeconomic influences on migration behavior, e.g., changes in the social welfare system or changes in the level of unemployment rates (Decressin, 1994) and ζ_{aht} is the error term.[13]

Several authors suggest different measures of labor market tightness in specification of equation (6). Jackman/Savouri (1992) use vacancy rates in addition to unemployment rates, Juarez (2000) uses employment growth or employment rates, and Fields (1979) favors unemployment rates. Unfortunately comparable data for all countries are available for employment rates (i.e. employment as a share of resident active age population), only. Thus we focus on this measure of labor market tightness. Finally, as a proxy for variability of GDP per capita we follow Bentiviogli/Pagano (1999) and use the standard deviation of GDP per capita over the last three years.[14] Also we were unable to secure data on these variables for all countries for the complete time period. In particular we have no data for the U.K and we miss data on GDP for the countries reporting on NUTS III level (i.e. Denmark, Estonia, Slovenia) before 1995. Furthermore for Italy and Spain we exclude the island NUTS II regions of Sicily, Sardinia and Canaries and the Baleares from estimation.[15]

{Table 6 Around here}

Table 6 displays the results of decomposing the standard deviation of these explanatory variables into a component due to the variance across sending-receiving region pairs (the between standard deviation) and into a component, due to variation across time (the within standard deviation). The first of these gives indication of the size of regional disparities in the respective countries. The table thus indicates that both regional GDP per capita and employment rate disparities in the candidate countries are by and large comparable to those in most EU member states.

{Table 7: Around Here}

Table 7 presents the results of the regressions. It suggests that gross migration rates respond moderately to economic variables in the current EU member states. For most of the EU countries analyzed (all but Italy and Belgium) we find a significant or at least marginally significant impact of regional per capita income disparities on migration. Furthermore, for some of the countries (Italy, Belgium and Spain) the coefficients on employment rate disparities are significant or on the verge of significance. Coefficients on the differences in variability of GDP by contrast attain significance in the case of the Netherlands only. This suggests that in contrast to the more distant migration analyzed in Bentivogli/Pagano (1999) differences in risk aversion play only a minor role in the migration decision for migration within a country.

For the candidate countries, we find that per capita GDP differences are significant and of the expected sign for Estonia, only. They are significant but have an unexpected sign for Hungary - suggesting that migrants move from high income to low income regions in this country. For all other countries GDP differences remain insignificant. Furthermore, differences in employment rates are insignificant for only two countries (Hungary and Poland). These results thus suggest that migration in the candidate countries is somewhat less responsive to regional income disparities than in EU member states.

The most robust result for both candidate countries and EU – member states is, however, that bilateral fixed effects explain the majority of the variation in gross place to place migration. R^2 values after including GDP differentials, employment rate differentials and differences in variation in GDP mostly increase by 1 to 2 percentage points relative to a specification with only bilateral fixed effects. Only for Estonia and Austria does the inclusion of measures of regional disparities increase the explanatory power of our estimates. This suggests that a substantial part of gross migration in both the EU and candidate countries is driven by factors other than economic motives.[16]

For this reason and since most models, which consider migration as an equilibration mechanism in the face of regional disparities focus on net migration, we estimate equation (6) using net rather than gross migration rates as dependant variable.[17] Results of this specification (Table 8) reconfirm much of the previous findings. In particular net migration in most EU-member states is significantly correlated with regional per capita GDP disparities but insignificantly correlated with these disparities in candidate countries. In Poland and Belgium furthermore we get significant coefficients with an unexpected sign. Differences in the variation of GDP are also insignificant in both EU and candidate countries.

Focusing on net migration, however, increases the significance of employment rate differentials in a number of EU – member states (Belgium, Spain and the Netherlands), while correlations of net migration with employment rate disparities in the candidate countries remain insignificant in all cases but Slovenia. Furthermore, marginal effects of regional GDP disparities increase when significant; suggesting that net migration is more strongly correlated with regional GDP disparities than gross migration rates. This is also reconfirmed when considering the additional explicative power of regional disparity measures in explaining net migration rates. The increases in R^2 values relative to a specification with only fixed effects are more sizeable than in the case when gross migration is the dependent variable.

{Table 8 around here}

3.1 A Decomposition

In summary the results presented in tables 7 and 8 imply that migration is less responsive to regional disparities in candidate countries than in most member states, where the most important difference is the lower responsiveness of candidate country migration to disparities in per capita GDP levels. To quantify the effect of these differences on migration in the candidate countries relative to the EU we perform a decomposition, in which we estimate the increase in migration that would occur if responsiveness of migration to regional disparities were as high as in an EU country in one of the EU member states.

Formally, this can be done by denoting a and b as estimates of the coefficients on income and wage disparities in a particular member state. The relative increase in total migration in the candidate country (ΔM) under the assumption that the responsiveness to wage and income disparities were equal to that in the member states, while all other factors remain equal, would then be given by:

$$\Delta M = \frac{\sum_a \sum_h \left(e^{a\ln(\mu_{at}-\mu_{ht})-b\ln(u_{at}+u_{ht})+c\ln(\sigma_a^{\;2}-\sigma_h^2)-\sum_a \sum_{h\neq a}\phi_{ah}+\sum_t \tau_t+\varsigma_{aht}} - M_{aht}\right)}{\sum_a \sum_h M_{aht}} \qquad (7)$$

where c, ϕ, and τ are the parameters estimated from equation (6) for the candidate country.

{Table 9 Around here}

We perform this calculation for both net and gross migration using Spain, Italy and the Netherlands as baseline EU member states.[18] Results (in table 9) suggest that the lower responsiveness of migration to regional disparities in the candidate countries contributes to low internal migration. For most countries our calculations increases in gross migration should be between 10% to 50% if the reaction of migration to regional disparities were similar to Spain, Italy or the Netherlands. Extreme increases are indicated throughout for the Czech Republic, where migration should increase by a factor of between 2 and 5. Slovene gross migration seems to already have converged to the levels of these countries. When focusing on net migration, however, our calculations suggest that migration figures should more than double to reach western European level in almost all candidate countries and should multiply by a factor of five to ten in a number of instances.

4 Conclusions

This paper used data on inter-regional migration for 9 current EU – member states and 7 countries that will join the European Union in 2004 or are negotiating on membership, to compare regional migration patterns in these countries. Our most important results are first, that interregional migration is low by EU standards in candidate countries and has been falling throughout the 1990s and second, that the responsiveness of migration to regional disparities is substantially lower in the member states than in the EU. We predict that in the typical candidate country gross migration should increase by between 10 to 50% if the responsiveness of gross migration to regional disparities were comparable to the member states and increases net migration should range between a factor of 2 and more than 10.

The findings thus suggest that low migration rates are one of the major obstacles to equalization of regional disparities as well as to effective absorption of asymmetric shocks in the candidate countries. On the policy side this clearly suggests that policies designed to reduce barriers to migration in the candidate countries should have a high priority. Unfortunately we are unable to answer the question, why the responsiveness of migration is so low in the candidate countries, which could provide orientation as to which policies could be most helpful in increasing migration.

We would, however, argue that a policy framework to address the low internal migration rates in candidate countries should take a relatively broad view on migration and should encompass a multitude of factors such as housing and capital market imperfections (to overcome liquidity constraints), improving spatial matching and reviewing labor market institutions (in particular employment protection regulation). Clearly, for policy purposes it would be interesting to know which of these factors would be most effective in increasing the willingness to migrate. This, however, is beyond the evidence presented in this paper and must be left to future research.

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Table 1: Data Sets used Countries, time periods and nature of the data

	Regional Disagregation	Number of Regions	Average Size*	Years Available	Place to place
Austria	NUTS II	9	898.1	1996-1999	yes
Belgium	NUTS II	11	928.5	1990-1999	yes
Germany	NUTS I	16	5127.3	1990-1993	yes
Denmark	NUTS III	15	354.6	1990-1999	yes
Spain	NUTS II	17	2316.6	1990-1999	yes
Italy	NUTS II	19	2983.3	1990-1996	yes
Netherlands	NUTS II	12	1313.4	1990-1999	yes
Sweden	NUTS II	6	1048.8	1990-1999	yes
U.K	NUTS I	12	4947.5	1990-1996	yes
Czech Republic	NUTS II	8	1286.2	1992-1999	yes
Estonia	NUTS III	5	275.8	1990-1999	yes
Hungary	NUTS II	7	1441.7	1990-1999	yes
Poland	NUTS II	16	2415.8	1990, 1995-1999	yes
Romania	NUTS II	8	2811.1	1994-1999	no
Slovenia	NUTS III	12	164.9	1991-1999	yes
Slovakia	NUTS II	4	1348.4	2000	no

Notes: NUTS=Nomenclature Unifie des Territoire Statistique , * in thousand inhabitants 1999, Source Eurostat New Cronos

Table 2: Migration indicators by country and year

	Gross MigrationRates[1)]		Net Migration Rates[2)]		Share of net Migration[3)]	
	1992	1999	1992	1999	1992	1999
Austria		0.93		0.054		5.79
Belgium	1.26	1.28	0.123	0.086	9.77	6.73
Germany	1.88	n.a.	0.152	n.a.	8.09	n.a.
Denmark	3.38	3.41	0.090	0.095	2.66	2.77
Spain	0.53	0.76	0.043	0.099	8.12	12.96
Italy	0.54	n.a.	0.097	n.a.	17.94	n.a.
Netherlands	1.63	1.69	0.079	0.063	4.85	3.75
Sweden	1.63	1.87	0.095	0.182	5.83	9.75
U.K	2.70	n.a.	0.132	n.a.	4.88	n.a.
Czech Republic	0.57	0.50	0.009	0.063	1.64	12.61
Estonia	0.87	0.53	0.203	0.024	23.24	4.64
Hungary	1.49	1.32	0.094	0.054	6.30	4.11
Poland[a)]	0.37	0.29	0.053	0.033	14.48	11.20
Romania	n.a.	1.23	n.a.	0.013	n.a.	1.09
Slovenia	n.a.	0.30	n.a.	0.021	n.a.	7.15
Slovakia[b)]	n.a.	0.22	n.a.	0.023	n.a.	10.25

Notes: Gross and net migration rates are measured in % of the population. a) Polish data for 1992 are 1990 figures b) Slovak data are from the year 2000. n.a. – data not available. 1) Figures are in %, see equation 1 for a definition of net migration flows. 2) Figures are in %, see equation 2 for a definition of net migration flows. 3) Figures are in %, see equation 4 for a definition of the share of net migration flows. Source: Eurostat New Cronos.

Table 3 Migration by Demographic Characteristics of Migrants

	share of females in total internal migration		share active of active aged in total internal migration	
	1992	1999	1992	1999
Austria	n.a.	47.42	n.a.	74.79
Belgium	50.25	49.81	70.25	70.51
Denmark	47.88	48.10	74.89	76.78
Spain	49.61	48.44	63.97	70.66
Italy	46.89	n.a.	68.92	n.a.
Netherlands	49.21	49.18	67.34	71.21
Sweden	49.70	51.06	68.77	77.76
U.K	51.72	n.a.	63.33	n.a.
Czech Republic	n.a.	52.42	n.a.	64.49
Estonia	52.42	58.21	52.01	57.69
Hungary	49.98	53.33	62.80	66.41
Romania	n.a.	56.01	n.a.	74.22
Slovenia	n.a.	55.86	n.a.	n.a.
Slovakia*)	n.a.	54.12	n.a.	68.47

Notes: Figures are percentages of total migrants *) Slovak data are from the year 2000. n.a. – data not available, Source Eurostat New Cronos

Table 4: Share of moves between neighboring regions and intertemporal correlations of migration rates

	Share of Migration Flows among neighbor Regions[a)]		Share of neighbor relationships [b)]	Correlation[b)] 1992-1999
	1992	1999	n.a.	n.a.
Austria		66.3	23.4	n.a.
Belgium	64.2	66.5	26.7	0.79
Denmark	53.4	52.2	17.3	0.84
Germany	53.4	n.a.	19.2	n.a.
Spain	36.6	37.5	17.6	0.51
Netherlands	60.8	60.0	25.8	0.92
Italy	28.7	n.a.	14.5	0.80
Sweden	48.1	55.9	26.3	0.48
Czech Republic	63.6	65.2	30.0	0.55
Estonia	71.1	72.6	60.0	0.62
Hungary	n.a.	77.2	34.4	n.a.
Poland	58.4	62.3	22.6	0.71
Slovenia	65.8	64.5	37.8	0.64

Notes: a) Columns report the share of total migration among neighboring regions as a percentage of total migration flows in 1999 and 1992, respectively; b) column reports the share of neighbor relationships in a country this is calculated by observing that in a country with n regions there are n*(n-1) pairs of sending and receiving regions. If m of these region pairs are contingent the share of contingency relationships in the total number of sending and receiving region pairs is given by $s = m / n(n-1)$. c) Column reports the correlation coefficient (across regions) of net emigration in % of population between 1992 and 1999. n.a. - data not available. Source: Eurostat New Cronos.

Table 5: External Migration in % of resident population

	Net Migration Abroad[a)]		Gross Emigration abroad [b)]	
	1992	1999	1992	1999
Austria	n.a.	n.a.	n.a.	0.9343
Belgium	n.a.	0.2659	n.a.	0.4044
Denmark	0.2216	0.1672	0.6172	0.7772
Germany	0.9742	n.a.	0.8971	n.a.
Spain	0.0948	0.3225	0.0052	0.0042
Italy	0.0993	n.a.	0.1001	n.a.
Netherlands	0.3068	0.3815	0.3184	0.3745
Sweden	0.2467	0.1797	0.3071	0.4126
Czech Republic	0.0853	n.a.	n.a.	0.5088
Estonia	-2.1756	-0.0447	2.4038	0.1475
Hungary	0.1113	0.1753	0.0425	0.0244

Notes: a) columns report net immigration (immigration – emigration) abroad in % of total population b) columns report gross emigration abroad in % of total population. n.a. - data not available. Source: Eurostat New Cronos

Table 6: Standard deviations of independent variables

	Differences in per capita GDP		Differences in employment rates		Differences in Variability of GDP	
	between	within	between	within	between	within
Austria	0.293	0.012	0.192	0.009	0.648	0.757
Belgium	0.414	0.025	0.318	0.020	1.099	1.254
Denmark	0.296	0.023	0.194	0.014	0.765	0.645
Germany	0.588	0.097	0.152	0.041	0.712	0.701
Spain	0.286	0.018	0.149	0.026	0.646	1.290
Netherlands	0.214	0.039	0.081	0.024	1.186	1.509
Italy	0.370	0.019	0.199	0.023	0.724	1.524
Sweden	0.146	0.028	0.052	0.012	0.624	1.203
Czech Republic	0.372	0.046	0.059	0.016	0.997	0.820
Estonia	0.468	0.042	0.087	0.024	1.397	0.444
Hungary	0.322	0.080	0.276	0.024	1.967	1.250
Poland	0.233	0.042	0.156	0.005	1.421	0.962
Slovenia	0.196	0.031	0.170	0.024	0.811	0.490

Note: Table reports within and between components of standard deviations. Source: Euostat New Cronos, Cambridge Econometrics

Table 7: Estimation Results of Equation (6) dependent variable Gross Migration

	GDP Differences	Employment rate Differences	Differences in variability of GDP	$R^{2b)}$ (NOBS)	R^2 only dummies
Austria	-5.593**	2.535	0.021	0.66	0.60
1996-1999	(2.896)	(1.887)	(0.052)	(288)	
Belgium	0.794	-0.656*	0.007	0.82	0.81
1993-1999	(0.477)	(0.391)	(0.019)	(770)	
Denmark	-0.658**	0.122	0.0001	0.89	0.87
1995-1999	(0.302)	(0.532)	(0.0003)	(1050)	
Germany	-1.406***	1.144	0.036	0.90	0.89
1990-1993	(0.376)	(0.860)	(0.032)	(460)	
Spain[a)]	-0.993**	-0.414*	0.001	0.98	0.96
1990-1999	(0.173)	(0.251)	(0.004)	(1890)	
Netherlands	-2.587***	-0.193	0.021**	0.80	0.78
1990-1999	(0.305)	(0.148)	(0.009)	(1188)	
Italy[a)]	-0.150	-0.883***	0.001	0.91	0.90
1990-1996	(0.342)	(0.157)	(0.006)	(1628)	
Sweden	-4.513***	-0.261	-0.005	0.89	0.87
1991-1990	(0.870)	(0.269)	(0.013)	(348)	
Czech Republic	3.078**	-0.174	-0.026	0.68	0.66
1993-1999	(1.156)	(0.167)	(0.0267)	(392)	
Estonia	-1.310**	3.283	0.031	0.79	0.65
1990-1999	(0.481)	(2.184)	(0.105)	(80)	
Hungary	0.464***	-0.702***	0.001	0.94	0.89
1990-1999	(0.113)	(0.051)	(0.008)	(336)	
Poland	0.020	-0.492***	0.004	0.92	0.91
1995-1999	(0.160)	(0.126)	(0.009)	(1200)	
Slovenia	-0.808	0.590	-0.106	0.73	0.73
1995-1999	(1.088)	(1.058)	(0.111)	(341)	

Notes: Dependent variable: gross migration rates in % of the population. a) Estimates for Italy and Spain exclude the islands Canaries, Baleares, Sicilly and Sardinia, *** (**) (*) signify significance at the 1% (5%) and (10%) level respectively. Values in brackets are standard errors of the estimate. b) Values in brackets are Numbers of Observations (NOBS)

Table 8: Estimation Results of Equation (1) dependent variable Net Migration

	GDP Differences	Employment rate Differences	Differences in variability of GDP	$R^{2\,b)}$ NOBS	R^2 only dummies
Austria	-13.744**	7.788	0.034	0.78	0.75
1996-1999	(6.190)	(4.352)	(0.145)	(143)	
Belgium	5.239**	-6.221***	-0.004	0.77	0.69
1993-1999	(2.645)	(2.003)	(0.039)	(380)	
Denmark	0.983	-2.656	0.000	0.70	0.69
1995-1999	(1.101)	(1.918)	(0.001)	(522)	
Germany	-3.367**	3.897	0.225***	0.81	0.75
1990-1993	(1.097)	(2.411)	(0.092)	(230)	
Spain[a)]	-4.677***	-5.872***	-0.005	0.75	0.66
1990-1999	(1.221)	(1.792)	(0.025)	(938)	
Netherlands	-4.210***	-0.961**	0.009	0.53	0.49
1990-1999	(1.005)	(0.445)	(0.029)	(592)	
Italy[a)]	-5.994***	-0.200	-0.016	0.80	0.75
1990-1996	(1.125)	(0.458)	(0.018)	(814)	
Sweden	-2.486	-1.512	-0.026	0.75	0.70
1991-1990	(3.076)	(1.000)	(0.055)	(174)	
Czech Republic	4.072	-0.696	-0.187**	0.81	0.61
1993-1999	(3.410)	(0.464)	(0.081)	(385)	
Estonia	-3.019	8.440	0.144	0.60	0.34
1990-1999	(2.067)	(9.325)	(0.483)	(40)	
Hungary	3.645	1.134	0.043	0.71	0.47
1990-1999	(2.543)	(0.926)	(0.045)	(168)	
Poland	1.758***	0.438	-0.009	0.73	0.62
1995-1999	(0.571)	(0.490)	(0.036)	(589)	
Slovenia	6.646	-5.417**	-0.408	0.61	
1995-1999	(2.454)	(2.661)	(0.315)	(149)	

Notes: Dependent variable: net migration rates in % of the population. a) Estimates for Italy and Spain exclude the islands Acores, Baleares, Sicilly and Sardinia, *** (**) (*) signify significance at the 1% (5%) and (10%) level respectively. Values in brackets are standard errors of the estimate. b) Values in brackets are Numbers of Observations (NOBS)

Table 9: Results of a decomposition of migration flows

	Italian coefficients	Spanish coefficients	Dutch coefficients
		Gross Migration	
Czech Republic	315.7	565.3	212.4
Estonia	118.6	147.1	327.4
Hungary	99.8	101.8	130.4
Poland	154.8	116.1	105.3
Slovenia	99.0	98.8	103.6
		Net Migration	
Czech Republic	500.37	260.30	1326.74
Estonia	339.70	554.93	982.34
Hungary	374.90	306.57	174.51
Poland	168.83	159.21	470.23
Slovenia	210.97	594.29	158.09

Note: Table reports the estimated migration (in % of migration in the last year of observation) if migration were as responsive to regional disparities as in the Netherlands, Italy and Spain, respectively. See equation (7) for a formal definition.

NOTES

[1] Financial support from the European Commission 5th Framework Programme – ACCESSLAB project is gratefully acknowledged. The author thanks the participants of the WIIW International Economics Seminar and the WIFO Research Seminar for helpful comments. Mistakes remain the responsibility of the author.

[2] For example Fidrmuc/Horvath/Fidrmuc (1999) argue that lacking regional mobility was one of the economic causes for disintegration of Czechoslovakia.

[3] For a number of EU member states data are available back to the 1970's. We limit our analysis to the 1990s to provide for similar time periods for both current EU member states and candidate countries.

[4] We performed similar analysis as below for other years as well as for data at different regional aggregations in earlier versions of this paper. The results of this analysis are comparable to those presented below and are available from the author.

[5] Furthermore, in Poland data for the year 1990 are not place to place data and the breakdown by age groups and gender presented below is also not available on a place to place basis.

[6] Division by two is necessary to avoid double counting since each outflow for one region is also an inflow for another region.

[7] These churning flows can be explained heterogeneity, either of individual tastes and characteristics or regional demand for labour (Fields, 1979), or through different life-cycle positions of individuals (e.g. students migrating to their place of education). Mueser (1997) shows that churning may also occur among ex-ante homogenous individuals due to endogenous wealth effects arising, for instance, from land prices increases due to exogenous shocks. Finally, spatial search models (Jackmann/Savouri, 1990, Molho, 2000, Juarez, 2000) predict that churning may result from stochastic matching, if workers do not search exclusively in their region of residence.

[8] Interestingly the increase in net migration in the Czech Republic is primarily due to the increase in migration from Prague to its environs (see: Fidrmuc/Huber, 2003).

[9] Unfortunately, the data on age and gender of migrants is not available on a place to place basis.

[10] This is calculated by observing that in a country with n regions there are n*(n-1) sending and receiving region pairs (since migration within the region is not measured). If m of these pairs are contingent, the share of contingency relationships in the total number of sending and receiving region pairs is given by $s = \frac{m}{n(n-1)}$.

[11] Furthermore, the limited evidence available suggests that long distance moves declined more strongly in candidate countries between 1992 and 1999. In both Hungary and the Czech Republic moves covering a distance of more than 100km were 18% below their 1992 level, moves covering a distance of less than 100km were 10% below the 1990 level.

[12] This is also owed to restrictive immigration regulations in EU member states, which are the primary destination countries for candidate countries emigrants.

[13] We give preference to a bilateral fixed effects specification over a specification with sending and receiving region fixed effects, because the later may be considered a restricted version of the former (Hui/Wall, 2001) and because information criteria such as the Akaike information criterion suggest that inclusion of bilateral fixed effects improves the model fit substantially.

[14] We use the previous two years when three lags are unavailable.

[15] Data on employment rates and GDP per capita for the NUTS I and NUTS II regions were provided by Cambridge Econometrics, for the NUTS III regions of (Denmark, Estonia, and Slovenia) this data was taken from the Eurstat Cronos database.

[16] We performed a number of robustness checks for this regression. In particular we excluded the differences in GDP variability, and experimented with specifications including distance between sending and receiving regions, as well as lagged variables to reduce potential endogeneity. None of this changes the qualitative results.

[17] Note that in this case we loose half of the observations since net migration is equal (but oppositely signed) between any pair of sending and receiving regions.

[18] This choice was guided by an attempt to use countries both from the north of the EU, with relatively low aggregate unemployment rates and higher labour market flexibility and from the South, where unemployment rates are somewhat higher and labour market flexibility is lower.

Research Report Prepared for the ACCESSLAB Project

Commuting among Low-skilled Workers in Hungary*

Tamás Bartus**

ABSTRACT

Although the unemployment rate is decreasing in Hungary during the last ten years, it is still high in those villages where it was the highest (above 20 percent) in the mid 1990s. It was suggested that the persistence of rural unemployment is due to the relatively high costs of commuting. This paper addresses the question of how commuting behavior is influenced by the distance between place of residence and place of work. The question is examined using retrospective information taken from a survey conducted among unemployed. The findings are as follows. (1) Commuters receive relatively high wages, and afford relatively long commuting, provided that travel expenses are covered. (2) The difference in wages between commuters and stayers remain after adjusting for several wage determinants. This means that the wage difference between commuters and stayers is a compensating wage differential. (3) The association between commuting and the coverage of travel expenses is very strong. (4) Independently of coverage of travel expenses, women have shorter commutes than men. These findings indicate that commuting costs substantially constrain spatial labor mobility, especially that of women.

Keywords: Commuting, Spatial Mismatch Hypothesis, Compensating Wages

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1. INTRODUCTION

Although the unemployment rate is decreasing in Hungary during the last ten years, it is still high in those villages where it was the highest (above 20 percent) in the mid 1990s. In their earlier papers, János Köllő and Gábor Kertesi argued that persistent unemployment in villages is due to the fact that commuting costs substantially exceed the returns to commuting in terms of wages (Köllő, 1997; Kertesi, 2000). In other words, commuters do not receive compensating wages (Leigh 1986) for the direct monetary expenses and the time spent on commuting. If urban firms do not pay compensating wages for commuters, then residents of villages far from urban centers will suffer from high and persistent unemployment. This line of argument is similar to the well-known spatial mismatch hypothesis, which claims that the suburbanization of job opportunities accounts for the high unemployment rate among black inner-city residents (Kain 1992, Ihlanfeldt and Sjoquist 1998).

Few attempts were made to test the above mentioned explanation in the Hungarian context. Köllő (1997) constructed a transportation database with settlements as units of observation. Using this database he showed that if there are no public transportation links, commuting with cars would use up a substantial part of the expected wages. Public transportation links are especially underdeveloped in regions where villages with high unemployment rates are typically situated. The transportation database also contains lower-bound estimates of travel expenses.[1] Kertesi (2000) relied on these estimates when analyzing the 1996 micro-census of the Hungarian Statistical Office. He found that the probability of commuting decreased with commuting costs, which was measured indirectly, as the difference between the unemployment rates of the place of residence and those urban centers that could be reached at the expense of 4,000 Forint. He also found that low-skilled villagers were more severely constrained in commuting by transport costs than were their high-educated counterparts – a finding that motivates us to restrict the forthcoming analysis to low-skilled workers (those without college or university diploma). Unfortunately, the actual commuting costs are not observed in these studies.

The purpose of this paper is to test the hypothesis of commuting costs using individual-level data that also contain information on the actual costs of commuting. More specifically, this paper attempts to answer the following questions: (1) How do wages and the frequency of commuting depend on travel distance and commuting cost?; (2) Are compensating wages paid

[1] It is assumed that the median villager takes a train or a bus if these are available, and considers driving if and only if the urban centres are not accessible by means of public transport, within reasonable time limits.

for costly commuting? (3) What is the relative effect of wages and commuting costs on the probability of commuting? (4) Are they gender differences in the compensating wage effect? The last question is motivated by the finding of previous research that women have shorter commutes than men (Cooke and Ross 1999).

2. DATA AND VARIABLES

Our analyses are based on a survey that took place among unemployed people who were entitled to unemployment benefits and got a job in the period between 18 of March and 7 of April 2001 (N=105,924). In this period 9474 people got a job, out of which 8339 people completed the questionnaire (Köllő, 2002). The questionnaire contains both retrospective questions about the previous job and questions about the new job. Information covers the characteristics of job and the firm, the names of the settlement where the job is located, place of residence, and commuting time.

In principle, the availability of information about two jobs for each respondent offers the opportunity to double the sample size or study the relationship between changes in wages and changes in commuting distance (Leigh 1986). However, this study makes use of retrospective information. This is due to the fact that only the retrospective questions are free of two important data problems that characterize information about the new job. First, information about travel expenses is not available concerning the new job because the relevant part of the questionnaire contains an error. Second, the questionnaire assumes that respondents do not know the exact value of their prospective salaries. They were therefore asked to provide an estimate of the new salary in terms of a minimum and a maximum value. Although the wage could be measured as the mean of the two estimated values, this measure would not be reliable because the difference between the two values is substantially different in a considerable proportion of cases. The disadvantage of relying on retrospective information is that reported values are subject to recall biases.

Table 1 lists the variables used in this study and the definitions thereof. The last monthly wage variable is the gross monthly salary in the last month before loosing the job, recorded in thousands of Hungarian Forint. Commuting is a dummy variable that takes value 1 if the place of residence and place of work are different and 0 otherwise. Commuting distance is the distance between place of work and place of residence as measured on public roads. Commuting distances were matched to our data from a unique database containing the distance matrix of Hungarian settlements. Since there are 3157 settlements, the database contains 3157^2=9,966,649 (almost ten millions) observations and three variables (the codes of two settlements and the distance between these settlements). (Note that the distance between two settlements occurs twice in the database.) The operational definition of distance is distance between the centers of the settlements measured on the shortest available public road. Unfortunately, the distance figures do not measure the actual travel distances of workers. Of particular importance is the fact that the distance of a settlement from itself is zero, thus people who work in their place of residence are assumed to have zero commuting distance. The original values higher than zero were transformed into five categories (10,20,30,40,50) using the 10*int*((*d-1)*/10)+10 transformation, where *d* is the original value and *int*(*d*) returns the integer of *d*.

TABLE 1 ABOUT HERE

Unfortunately, respondents were not asked to report on the actual value of money they spend on travel. Instead, they were asked to report on the employer's coverage of travel expenses. Thus, commuting costs are captured by the travel expenses variable. It takes the value 0 if the employer does not cover travel expenses, while it takes the value 1 if the firm covers a part or the full amount of travel expenses or it organizes the travel of workers at its own expenses.

Besides these variables, our analyses will control for other wage determinants like human capital, firm level characteristics and local unemployment rates. Human capital is captured by gender, education, and age. Education is a dummy variable taking the value 1 if the respondent has A-level and 0 if the respondent has only vocational or (some) elementary education. The type of occupation variable takes the value 1 if the respondent has a white-collar job and 0 if the respondent works in a manual occupation. Firm size records the number of employees at the firm. Local unemployment rate is the ratio of the number of unemployed to the number of economically active population within the micro-regions where the workplace is situated. Information about unemployment, economic activity and micro-regions are taken from the TSTAR 2000 database of the Hungarian Statistical Office. The TSTAR databases have settlements as observations and covers information about several economic, social and demographic variables. Finally, we will also control for the effect of the minimum wage. The minimum wage was substantially raised in 2001. The year variable is a dummy which takes the value 1 if the last monthly wage variable is observed in 2001 and 0 if the year of observation is 2000 (see also below).

Survey data are rarely free of data problems. We deleted those cases in which settlement codes or values of variables were nonsensical. The sample size was further reduced by deliberate decisions. A theoretically motivated decision was to exclude those unemployed who changed their place of residence during their unemployment spell. The reason is that migration might disturb the empirical relationship between commuting distance and commuting decisions (Ihlanfeldt and Sjoquist, 1998). In order to minimize the occurrence of outlier data points that would have an enormous effect on regression coefficients we deleted observations where (1) wages are higher than 100 thousands Forint; (2) commuting distance exceeds 50 km; (3) college or university education is reported; or (4) work is not carried out under a regular employment relationship (that is, part-time work, traineeship or no employment relationship characterizes the work situation). Criteria (3) and (4) guarantee that, consistent with our purposes described in the Introduction, we study a labor market segment in which people with relatively low education are matched to regular jobs. In order to minimize recall biases, we deleted cases where last wage data are prior to 2000. Finally, we deleted observations where any of our variables have missing values. These observations would be automatically excluded in regression analyses. As a result of these decisions, we are left with a sample of size 4599 for further empirical analyses. I will refer to this sample as the estimation sample throughout this paper.

3. THE EMPIRICAL MODEL OF COMMUTING

The hypothesis of commuting costs can be summarized as follows. Suppose an unemployed receives two job offers. One of the jobs is located in the current place of residence, the other job is located in another settlement at distance d from the place of residence. The unemployed prefers commuting if the value of the latter wage offer (w_d) minus the costs of commuting (c_d) is higher than the value of the local wage offer (w_0). (Note that standard value of time models imply that the full cost of commuting is the sum of the monetary costs and the costs associated with travel time (Fujita 1989, Brueckner, Thisse and Zenou 2002)). Otherwise the unemployed prefers to work in his or her place of residence (stayer).

The hypothesis of commuting costs implies that for any distance d,

$$(1) \qquad w_0 > w_d - c_d .$$

The first objective of our empirical analyses is therefore the assessment of this relationship. A special and important interpretation of equation (1) is that it expresses a partial relationship. This means that the wage difference (w_d-w_0) is a wage differential that compensates for costs associated with commuting (Leigh 1986), and is unaffected by possible differences in the composition of individual-level and firm-level characteristics between commuters and stayers.

In short, equation (1) expresses a claim about the relative amount of compensating wages. The compensating wage approach suggest the study of wages as a function of commuting distance and commuting costs.

Besides studying the spatial distribution of wages, the hypothesis of commuting costs aims at explaining the occurrences of commutes as a function of wages and commuting costs in order to explain the frequency of commuting as a function of commuting distance d. Since the unemployed prefers commuting if the value of the latter wage offer (w_d) minus the costs of commuting (c_d) is higher than the value of the local wage offer (w_0), the starting point of the analysis is the equation

$$(2) \qquad \Pr(I = 1) = F(w - c_d),$$

where I is a binary variable measuring commuting (I=1 for commuters, and I=0 for stayers).

It is reasonable to assume that the monetary cost of commuting is a linear function of distance. Let c be the monetary costs of traveling one km. Assume further that traveling has no fixed costs. Then equation (2) can be re-expressed as

$$(3) \qquad \Pr(I = 1) = F(w - cd).$$

Unfortunately, our data does not allow a direct estimation of equation (3). First, the monetary cost of traveling 1 km (c) is unknown. What we know is whether or not traveling involves monetary costs to be covered by the worker. Second, the measurement of commuting distance is not perfect. Due to the use of the distance matrix, people who work in their place of residence are assumed to travel 0 km. If d=0 for workers who do not have to travel to other settlements then equation (3) cannot be estimated using the standard statistical models for discrete choice problems, like the logit or the probit model.[2] Thus, we have the problem of not being able to

[2] This is due to technical reasons. Measurement creates a deterministic relationship between the absence of commuting and zero commuting distance. In probit and logit models, deterministic relationships are modeled with infinite parameter estimates, since in these models infinitely large coefficients guarantee that the occurrence of an event is one. Unfortunately, the convergence of the probit and logit models might be difficult to achieve if one of the coefficients is infinitely large. In order to secure the convergence of the iterative estimation, one should discard those observations in which the relationship between distance and commuting is deterministic. After deleting these observations, however, the sample will cover only commuters and thereby the model cannot be estimated.

estimate the effect of commuting distance on the probability of commuting. Therefore, we will estimate only equation (2). We will examine the relationship between distance and the chances of commuting using simple cross-tabulations.

4. EMPIRICAL ANALYSES

The empirical analyses proceed in three steps. First, we describe the distribution of commuting distances and the distribution of wages as a function of commuting distances. The descriptions will make use of simple tables. Second, we examine the relationship between wages and factors influencing commuting costs. Here we aim to answer the question whether compensating wage differentials are paid for costly commuting. Finally we examine both the direct effect of wages and commuting costs on the likelihood of commuting.

Before proceeding, it is useful to examine the data. *Table 2* shows the means of the variables used in subsequent analyses. An apparent characteristic of our sample is that men are overrepresented: the proportion of men to women is 3:1. Recall that our sample is taken from a survey conducted among those registered unemployed who found a job in a certain time period. The comparison of the estimation sample to the full dataset of unemployed revealed that the men/women ratio is higher in the estimation sample. Additionally, this comparison also revealed that women are more likely to have high, that is, general A-level education. (Notice the difference in the proportion of general A-level education between men and women in *Table 2*.) These findings suggest that the large men/women ratio is due to the fact that employers demand (male) workers with specialized skills rather than (female) workers with relatively general skills.

TABLE 2 ABOUT HERE

Commuting is not a rare phenomenon: the frequency of commuting is 40 percent among men and somewhat less, 34 percent among women. The average commuting distance 9 km among men and 6 km among women. In order to recover the gender-specific average commuting distances *among commuters*, these figures must be divided by the gender-specific probabilities of commuting (0.4 and 0.34, respectively). Thus, the average commuting distance among male and female commuters are 20 and 18.6 km, respectively. The majority of workers do not receive coverage of travel expenses. Again, in order to recover the gender-specific proportions *among commuters*, these figures must be divided by the same gender-specific probabilities. Note that men have a slight advantage of 6 percentage points over women in this respect. The average age in our sample is about 40 years (recall that the age variable takes the value zero if the respondent is of age 18). The sample is composed of people who have low education and who work in manual occupations, especially among men. An important characteristic of our sample is that only 2 percent of the cases are drawn from micro-regions where unemployment is higher than 15 percent. This means that our sample is not appropriate to study the situation of workers who live micro-regions with severe unemployment. The rare occurrence of such micro-regions should not be surprising. To repeat, our sample cannot contain information about those registered unemployed who did not get a job in the period between 18 of March and 7 of April 2001. It is likely that people living in micro-regions with extremely high unemployment were not able to get a job in this period. Finally, note that the majority of the respondents reported the last earnings data in 2000.

4.1. The Distribution of Commutes and Wages

In this subsection, we examine the distribution of commutes and the distribution of last monthly wages. *Table 3* displays these distributions by travel expenses, for both sexes separately (see panels *A* and *B*, respectively). The most striking finding is that commuting is very rare if the commuter does not receive coverage of his or her travel expenses from the employer. In the absence of such coverage, commuting distances are short. If travel expenses are not covered, no women and only four men are willing to travel at least 41 kilometers. The discouraging effect of commuting costs is easy to understand if we look at the corresponding wage figures. In the subsample of people receiving no coverage, even long-distance commuters earn the same as the stayers. Wages are not increasing with commuting distance, thus there are no incentives to engage in costly commuting.

TABLE 3 ABOUT HERE

Contrary to this, commuting occurs frequently if employers receive coverage of travel expenses from the employer. Similar to those who do not receive coverage of travel expenses, relatively long-term commuting distances are less frequent than short-distance commutes. The threshold distance above which commuting becomes rare is 40 km for men and 20 km for women. The fact that we observe frequent and longer commutes among workers receiving coverage of travel expenses is self-explanatory. Additionally, notice that wages increase monotonically with wages among both sexes. In short, not only the coverage of travel expenses but also the higher wages at distant workplaces explain the relatively frequent occurrence of long-term commuting.

These findings can be summarized as follows. Costs of commuting, measured with the travel expenses variable, have a substantial effect on the chances and the expected distance of commuting. As *Table 3* shows, the vast majority of commuters receive coverage of travel expenses. Additionally, commuters who bear all costs of commuting do not receive higher wages than stayers, while commuters who receive coverage of travel expenses earn more than stayers. Wages are an increasing function of commuting distance - but only among those who receive coverage of travel expenses. Thus, the two incentives for commuting, high wages and low commuting costs, are positively associated. In other words, high-wage firms attracting workers from other settlements are willing to cover parts of commuting costs, but low-wage firms attracting such workers are not willing to contribute to travel costs. Besides, similar to earlier studies, we found that women have shorter commutes than men. Independently of commuting costs, women mostly work relatively close (1-20 km) to their home. Contrary to this, we find a considerable proportion of men commuting more than 20 km, provided that travel expenses are covered.

4.2. Are Compensating Wages Paid for Commuting?

We proceed with the analysis of the relationship between wages and factors influencing commuting costs. Here we aim to answer the question whether compensating wage differentials are paid for costly commuting. To answer this question, we adjusted the raw difference in wages between commuters and stayers for various human capital, firm-level and regional characteristics using linear regression. For both sexes, two linear regression models were estimated. The only difference between these models is that only one of them includes travel expenses.

TABLE 4 ABOUT HERE

The estimates of the model are displayed in *Table 4*. First, consider the models where the interaction term between commuting and travel expenses is not included. The parameter estimate of commuting can be interpreted as a wage premium that compensates for commuting costs. The parameter estimates are statistically significant and positive for both sexes. Thus, commuters receive a compensating wage premium. This premium is about 3,500 Forint among men and 3,200 Forint among women.

Let us move to the models which also include the interaction term between commuting and travel expenses. Now the commuting variable should be interpreted as a compensating wage received by those who receive coverage of travel expenses, and the interaction term captures this wage premium among those who do not receive such coverage. The parameter estimates of commuting are statistically significant and positive for both sexes. Thus, we have reason to assume that commuters whose travel expenses are covered receive compensating wages. premium. The estimated size of the compensating wage is somewhat higher among men receiving coverage of travel expenses than among men who do not receive such coverage.

The interaction term has a significant parameter estimate among men, but is not significant among women. The sign of the interaction term is negative. The size of the estimate is somewhat larger than the estimate of the commuting dummy. However, the sum of the two coefficients is statistically not significant. This means that we have no grounds to believe that that commuters whose travel expenses are not covered receive compensating wages.

To sum, *Table 4* shows that compensating wages are received by those men and women who also receive coverage of their travel expenses. This finding is very similar to our previous finding that only commuters receiving coverage of travel expenses enjoy high wages. The regression analysis implies that the raw wage difference between commuters and stayers is not due to differences in the composition of relevant individual and firm-level characteristics between commuters and stayers.

Finally, looking at the parameter estimates of the other variables help us to assess the reliability of the results. If the parameter estimates of the other variables contradicted to theoretical expectations and previous empirical estimates, we could raise seriour doubts about the reliability of our results. Fortunately, the parameter estimates of the control variables are consistent with the estimates reported in earlier studies. The comparison of the constant terms indicates that men earn more than women. The coefficients of the other human capital variables (education, age and age-squared) have the expected signs. People working in white-collar occupations earn more than their counterparts working in manual jobs. The wages are also higher in larger firms. Unemployment rate has the expected negative effect on wages. Finally, the year dummy is positive, which is consistent with the increase in the minimum wage in 2001.

The analyses so far focused on the adjusted differences in wages between commuters and stayers. We proceed with estimating the compensating wage as a function of commuting distance. Similar to the previous analyses, we estimated two linear regressions, which differ only in the exclusion or the inclusion of the interaction between commuting distance and travel expenses.

TABLE 5 ABOUT HERE

Table 5 displays the estimation results. The parameter estimates of the commuting distance variable are statistically significant and positive for both sexes. Thus, the compensating wage is increasing with commuting distance. The interaction term has a significant parameter estimate among men, but is not significant among women. The sign of the interaction term is negative, and its magnitude is similar to that of the distance variable. This means that only workers receiving coverage of travel expenses enjoy compensating wages, and we have no grounds to believe that commuters whose travel expenses are not covered receive compensating wages.

Keeping in mind that commuting distance is recorded in 10 km units, the coefficients show that 10 km increase in commuting distance is compensated by 2,600 Forint among men and 2,000 Forint among women. This compensation scheme, however, holds only among those who receive coverage of travel expenses. Note that there is a gender difference in the amount of compensation. Men have an advantage of almost about 800 Forint among those who receive coverage of travel expenses.

The previous analyses assumed that an unit increase in commuting distance leads to a constant increase in wages. To check the assumption of constant effect, the distance variable was transformed into dummies. Then we replicated the previous analyses so that the single distance variable was replaced by the newly created dummies. A compact summary of the results are displayed in *Table 6*. The reader should keep in mind that the figures reproduced in the table are taken from linear regressions that also control for the previously used human capital, firm-level and micro-regional characteristics.

TABLE 6 ABOUT HERE

The results are mixed. One the one hand, the assumption of constant effect seems to hold among men receiving coverage of travel expenses, since the estimated difference between commuters with a given commuting distance and the stayers gets larger as the commuting distance increases. Additionally, no significant wage differences are found between any groups of commuters and stayers. On the other hand, wages do not increase monotonically with commuting distance among women. Women commuting 11-20 kilometers have a wage advantage over female stayers, regardless of travel expenses. Also commuting 31-40 kilometers in the presence of coverage of travel expenses guarantees a wage premium. But other groups of female commuters do not receive a wage premium. This finding is consistent with our previous finding that women rarely travel more than 20 kilometer to work.

4.3. Analysis of Commuting Decisions

It was found in the previous subsection that commuters receive a compensating wage, and this wage premium is increasing with distance among men. We also demonstrated that compensating wages are paid to those who also receive coverage of travel expenses. Earlier we also showed that commuting becomes rare with commuting distance, especially among those for whom commuting is costly. In this subsection, we aim to describe the partial effect of wages and travel expenses on the probability of commuting. The analyses reported in this subsection are equivalent with testing the hypothesis of commuting cost.

To test our hypothesis, we estimate a probit regression of commuting on wages, coverage of travel expenses and the other control variables. *Table 7* shows the estimation results. We expect that the wage and the coverage of travel expenses variables are positively associated with the probability of commuting. The signs of the parameter estimates of these two variables are consistent with our expectation. The parameter estimates are statistically significant. Thus, the probability of commuting increases with the wage offer, but it decreases if travel expenses are not covered. Apart from the coverage of travel expense variable, the variables have similar effects among both men and women. Note that unemployment in the place of residence has a positive, while unemployment in the micro-region has a negative effect on commuting.

TABLE 6 ABOUT HERE

Figure 1 shows the effect of the coverage of travel expenses on the probability of commuting. The panels show the predicted probabilities of commuting as a function of wages, separately for men and women. The upper line shows the predicted probabilities for those who receive contributions to travel costs, while the lower line shows the predicted probabilities for those who do not receive such contributions. For both sexes, the lines depict the following situation: the employee is 40 years old and has no A-level education, the local unemployment rate higher than 15 per cent, and the wage data is observed in year 2000. This situation is intended to model a situation which is closest to the situation of local labor markets with persistent unemployment.

FIGURE 1 ABOUT HERE

The figure clearly shows that coverage of travel expenses has a substantial partial effect on commuting decisions. If the travel expenses of a commuter are covered, then he or she commutes with an estimated probability of at least 80 percent. However, if all of the travel expenses must be paid by the worker, the predicted probabilities of commuting are much smaller. Predicted probabilities surprisingly slowly increase with the increase of last monthly wage. Thus, coverage of travel expenses has a large impact on commuting, and this effect is larger than the effect of wages.

5. CONCLUSIONS AND DISCUSSION

Although the unemployment rate is decreasing in Hungary during the last ten years, it is still high in those villages where it was the highest in the mid 1990s. The purpose of this paper is to test the hypothesis that of commuting costs. More specifically, this paper attempts to answer the following questions: (1) How do wages and the frequency of commuting depend on travel distance and commuting cost? (2) Are compensating wages paid for costly commuting? (3) What is the relative effect of wages and commuting costs on the probability of commuting? (4) Are they gender differences in the compensating wage effect?

The question is examined using retrospective information taken from a survey conducted among unemployed. The findings are as follows. Direct costs of commuting, measured with the travel expenses variable, have a substantial effect on the chances and the expected distance of commuting. The vast majority of commuters receive coverage of travel expenses. Additionally, commuters who bear all costs of commuting do not receive higher wages than stayers, while commuters who receive coverage of travel expenses earn more than stayers. Wages are an increasing function of commuting distance - but only among those who receive coverage of travel expenses. Thus, the two incentives for commuting, high wages and low commuting costs, are positively associated. In other words, high-wage firms attracting workers from other settlements are willing to cover parts of commuting costs, but low-wage firms attracting such workers are not willing to contribute to travel costs.

Regression analyses of wages showed that the difference in wages between commuters and stayers remain after adjusting for several wage determinants. Thus, the raw wage difference between commuters and stayers is a compensating wage differential, and not a wage difference that would reflect differences in the composition of relevant individual and firm-level characteristics between commuters and stayers. However, compensating wages are received by those men and women who also receive coverage of their travel expenses. A similar pattern was found when commuting was replaced by commuting distance.

On the basis of these findings, we can conclude that that commuting costs constrain labor mobility. This constraint is severe since commuting costs are negatively associated with wages. Regression analyses of commuting showed that commuting depends strongly on the coverage of travel expenses, and this effect is stronger than the positive effect of wages on commuting. The figure clearly shows that coverage of travel expenses has a substantial partial effect on commuting decisions. If the travel expenses of a commuter are covered, then he or she commutes with an estimated probability of at least 80 percent. However, if all of the travel expenses must be paid by the worker, the predicted probabilities of commuting are much smaller. Predicted probabilities surprisingly slowly increase with the increase of last monthly wage. Thus, coverage of travel expenses has a large impact on commuting, and this effect is larger than the effect of wages.

A consistent finding in the literature on commuting is that women have shorter commutes than men. The same difference was demonstrated in our study. Independently of coverage of travel expenses, women mostly work relatively close (1-20 km) to their home. Contrary to this, we find a considerable proportion of men commuting more than 20 km, provided that travel expenses are covered. This means that the absence of coverage of travel expenses constrain the commuting behavior of women stronger than that of men. Note that these are the women who are usually in a more disadvantaged labor market position. Our findings imply that the unwillingness of employers to cover the travel expenses of their workers is an additional cause of the disadvantaged position of women.

Our findings might suggest that coverage of travel expenses on the part of employers is a necessary condition for the reduction of persistent regional inequalities. This conclusion, however, neglects the possibility that employers will reduce labor demand as a reaction to increases in labor costs. If employers cut labor demand, it is difficult to predict the net effect of coverage of travel expenses on regional differences in unemployment rates. Knowing the precise effect of coverage of travel expenses on labor demand is a necessary condition for formulating firm policy recommendations on the basis of our empirical results.

A substantial limitation of our study is that our sample is probably not free of sample selection problems (Cooke and Ross 1999). Our sample stems from a survey of unemployed, and unsuccessful job searchers are not included in the sample. This might lead to the problem of self-selection if unobserved factors determining the success of job search (getting a job) are correlated with unobserved determinants of wages or commuting decisions. Fortunately, it is possible to make some comparisons between our estimation sample and the sample consisting of individuals who were not included in our analyses. The comparison of the means of the explanatory variables we used in the regression analyses between these two samples revealed no substantial differences with two exceptions. First, people with apprentice education are overrepresented, while people with a general A-level are underrepresented in the estimation sample. Second, as it was shown in *Table 2*, men are overrepresented in the estimation sample.

Note that choosing a general secondary school instead of an apprentice education is more popular among girls than among boys. Thus, the comparisons indicate that employers demand people with apprentice education instead of people having other kind of education, and these are the men who have the demanded type of education.

The reliance on a dataset of unemployed who got a job in a certain time period also raises several issues. As mentioned earlier, our sample cannot be considered as a representative sample of low-skilled Hungarian workers. Perhaps the most serious concern is raised by the fact that only 2 percent of the cases are drawn from micro-regions where unemployment is higher than 15 percent. This means that our sample is not appropriate to study the situation of workers who live micro-regions with severe unemployment. The rare occurrence of such micro-regions should not be surprising since our sample cannot include unemployed who did not get a job in a relatively short time period. It is likely that people living in micro-regions with extremely high unemployment were not able to get a job in this period. The use of a dataset that contains few observations from micro-regions with high unemployment is a serious limitation since the hypothesis of commuting cost aims to understand the persistence of high unemployment in these micro-regions. Nevertheless, the results from this study might be considered as an optimistic description of the situation of people residing in such micro-regions.

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TABLE 1
Definition of variables

Variable	*Definition and Notes*
Last monthly wage	Gross monthly salary in the last month before loosing the job, recorded in thousands of Hungarian Forint
Commuting	1 if the worker commutes; 0 otherwise
Commuting distance	Distance between place of work and place of residence as measured on public roads. Commuting distance is zero if the respondent does not commute. The original values higher than zero were transformed into five categories (10,20,30,40,50) using the 10*int*((*d-1)*/10)+10 transformation, where *d* is the original value and *int*(*d*) returns the integer of *d*.
Travel expenses	1 if the employer covers no part of travel expenses; 0 if the employer covers a part of travel expenses
Gender	1 if male; 0 if female
Education	1 if the respondent has A-level; 0 if the respondent has less education
Age	Age at the time of interview - 18
Age-squared	(Age at the time of interview – 40)^2
Type of occupation	1 if the respondent has a white-collar job 0 if the respondent works in a manual occupation
Firm size	Number of employees in the firm measured with three categories (1-5 employees, 5-50 employees, and 50< employees)
Local unemployment rate	Unemployment rate in the micro-region of the firm, measured with four categories (<5 %, 5-10 %, 10-15 %, 15< %)
Year of observation	1 if the last job was lost in 2000 0 if the last job was lost in 2001

TABLE 2
Means of the variables used in subsequent analyses

Variable	*All cases* (N=4599)	*Men* (N=3429)	*Women* (N=1170)
Last monthly wage	45.17	46.32	41.78
Commuting	0.39	0.40	0.34
Commuting distance	8.32	9.00	6.33
Travel expenses	0.60	0.59	0.65
Education	0.15	0.11	0.27
Age	20.31	20.52	19.71
Age-squared	110.48	112.73	103.89
Type of occupation	0.08	0.04	0.18
Firm size: 5-50 employees	0.47	0.48	0.43
Firm size: 50< employees	0.45	0.45	0.46
Local unemployment rate: 5-10 %	0.56	0.55	0.57
Local unemployment rate: 10-15 %	0.18	0.20	0.11
Local unemployment rate: 15< %	0.02	0.02	0.02
Year of observation	0.36	0.35	0.40

TABLE 3
Number of observations and means of last monthly wage by commuting distance and travel expenses

A) men

Commuting distance	*Travel expenses covered*		*Travel expenses not covered*	
	N	Mean	N	Mean
0	271	44.026	1785	44.094
1-10	348	45.339	131	42.000
11-20	339	48.811	65	38.785
21-30	210	54.452	22	45.455
31-40	177	57.729	10	52.100
41-50	67	67.313	4	43.500
Commuters	1,141	51.260	232	41.888
Total	1,412	49.872	2,017	43.840

B) women

Commuting distance	*Travel expenses covered*		*Travel expenses not covered*	
	N	Mean	N	Mean
0	63	44.222	707	39.352
1-10	152	43.434	30	33.133
11-20	123	48.203	17	48.941
21-30	39	48.436	5	36.600
31-40	20	52.500	3	39.667
41-50	11	61.182	0	
Commuters	345	46.791	55	38.691
Total	408	46.395	762	39.304

TABLE 4
Last monthly wage as a function of commuting and coverage of travel expenses: OLS parameter estimates

Independent variables	*Men*		*Women*	
Commuting	3.553	4.672	3.243	3.511
	(5.77)**	(7.07)**	(3.75)**	(3.82)**
Commuting * travel expenses		-5.493		-1.744
		(4.59)		(0.87)
Education	1.992	2.166	4.349	4.344
	(2.03)*	(2.21)*	(3.79)**	(3.79)**
Age	0.107	0.106	0.132	0.132
	(3.80)**	(3.75)**	(2.68)**	(2.68)**
Age squared	-0.010	-0.010	-0.002	-0.002
	(3.53)**	(3.58)**	(0.48)	(0.47)
Type of occupation	8.735	8.566	2.382	2.397
	(5.50)**	(5.41)**	(1.80)	(1.81)
Firm size: 5-50 employees	8.621	8.298	5.324	5.256
	(7.47)**	(7.20)**	(3.80)**	(3.74)**
Firm size: 50< employees	20.593	20.052	14.869	14.732
	(17.32)**	(16.83)**	(9.97)**	(9.82)**
Local unemployment rate: 5-10 %	-3.178	-3.127	-4.324	-4.307
	(4.43)**	(4.37)**	(4.68)**	(4.66)**
Local unemployment rate: 10-15 %	-5.886	-5.779	-3.876	-3.869
	(6.29)**	(6.19)**	(2.71)**	(2.71)**
Local unemployment rate: 15< %	-11.631	-10.896	-4.768	-4.724
	(6.16)**	(5.76)**	(1.54)	(1.52)
Year of observation	3.119	3.014	3.220	3.196
	(5.30)**	(5.13)**	(3.90)**	(3.87)**
Constant	32.071	32.425	25.117	25.185
	(13.84)**	(14.03)**	(10.59)**	(10.61)**
Observations	3429	3429	1170	1170
R-squared	0.23	0.23	0.28	0.28

Notes:
Absolute value of t statistics in parentheses
* significant at 5%; ** significant at 1%
Reference categories: no A-level; manual type of occupation; firm with <5 employees; local unemployment rate <5 percent; year of observation is 2000
The model also controls for industry using the following categories: agriculture, manufacturing and mining, construction, trade, transportation, service, and public administration

TABLE 5
Last monthly wage as a function of commuting distance and coverage of travel expenses: OLS parameter estimates

Independent variables	*Men*		*Women*	
Commuting distance	0.259	0.286	0.202	0.209
	(11.26)**	(12.04)**	(5.17)**	(5.10)**
Commuting distance * travel expenses		-0.260		-0.060
		(4.43)**		(0.57)
Education	2.215	2.411	4.231	4.240
	(2.29)*	(2.50)*	(3.71)**	(3.71)**
Age	0.114	0.111	0.139	0.140
	(4.11)**	(4.00)**	(2.84)**	(2.84)**
Age squared	-0.009	-0.009	-0.002	-0.002
	(3.29)**	(3.35)**	(0.53)	(0.51)
Type of occupation	8.851	8.665	2.620	2.624
	(5.65)**	(5.55)**	(1.98)*	(1.99)*
Firm size: 5-50 employees	8.128	7.891	5.486	5.449
	(7.14)**	(6.95)**	(3.94)**	(3.91)**
Firm size: 50< employees	19.260	18.847	14.676	14.606
	(16.33)**	(15.97)**	(9.89)**	(9.81)**
Local unemployment rate: 5-10 %	-3.115	-3.061	-4.452	-4.447
	(4.42)**	(4.35)**	(4.85)**	(4.85)**
Local unemployment rate: 10-15 %	-5.328	-5.250	-3.940	-3.941
	(5.80)**	(5.73)**	(2.78)**	(2.78)**
Local unemployment rate: 15< %	-10.759	-10.266	-4.788	-4.756
	(5.77)**	(5.51)**	(1.56)	(1.54)
Year of observation	2.750	2.640	3.252	3.234
	(4.72)**	(4.54)**	(3.96)**	(3.94)**
Constant	31.738	32.120	24.854	24.888
	(13.92)**	(14.12)**	(10.54)**	(10.55)**
Observations	3429	3429	1170	1170
R-squared	0.25	0.25	0.29	0.29

Notes:

Absolute value of t statistics in parentheses

* significant at 5%; ** significant at 1%

Reference categories: no A-level; manual type of occupation; firm with <5 employees; local unemployment rate <5 percent; year of observation is 2000

The model also controls for industry using the following categories: agriculture, manufacturing and mining, construction, trade, transportation, service, and public administration

TABLE 6

Adjusted differences in the average last monthly wage between commuters and stayers by commuting distance, travel expenses and gender

Commuting distance	*Men*		*Women*	
	Travel expenses covered	Travel expenses not covered	Travel expenses covered	Travel expenses not covered
Distance: 01-10 km	0.150	-0.472	1.209	-2.794
	(0.16)	(0.33)	(0.99)	(1.10)
Distance: 11-20 km	3.144	-3.647	4.799	10.147
	(3.22)**	(1.78)	(3.59)**	(3.03)**
Distance: 21-30 km	8.565	2.254	4.004	3.691
	(7.14)**	(0.66)	(1.76)	(0.54)
Distance: 31-40 km	11.331	5.847	7.519	-2.135
	(8.55)**	(1.15)	(2.42)*	(0.27)
Distance: 41-50 km	18.608	7.126	12.663	---
	(9.11)**	(0.89)	(3.03)**	---

Notes:

Adjusted differences are OLS parameter estimates of the linear regression of last monthly wage on categories of commuting distance, categories of commuting distance interacted with travel expenses, and various control variables. The control variables are: education, age, age-squared, type of occupation, firm size, industry, local unemployment rate and year of observation.

Absolute value of t statistics in parentheses;

* significant at 5%; ** significant at 1%

Reference categories: no A-level; manual type of occupation; firm with <5 employees; local unemployment rate <5 percent; year of observation is 2000

TABLE 7
Commuting as a function of last monthly wage and travel expenses: ML parameter estimates of a probit model

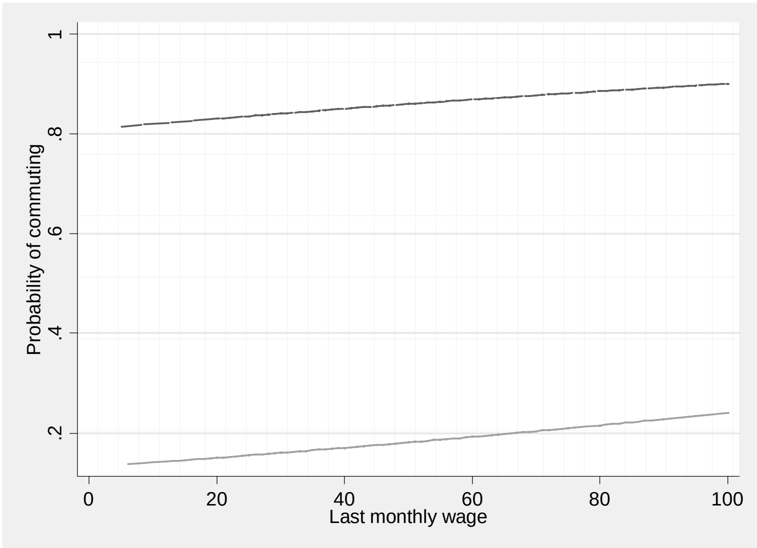
	Men	*Women*
Last monthly wage	0.004	0.002
	(2.77)**	(0.49)
Travel expenses	-1.991	-2.454
	(36.47)**	(23.51)**
Education	-0.074	-0.058
	(0.87)	(0.49)
Age	-0.010	-0.008
	(3.62)**	(1.32)
Age squared	-0.001	-0.000
	(2.09)*	(0.18)
Local unemployment rate: 5-10 %	-0.262	-0.095
	(3.98)**	(0.83)
Local unemployment rate: 10-15 %	-0.556	-0.295
	(6.38)**	(1.52)
Local unemployment rate: 15< %	0.037	-0.669
	(0.22)	(1.61)
Year of observation	-0.042	-0.005
	(0.74)	(0.04)
Constant	1.166	1.215
	(10.01)**	(4.95)**
Observations	3429	1170

Notes:
Absolute value of t statistics in parentheses
* significant at 5%; ** significant at 1%
Reference categories: no A-level; local unemployment rate <5 percent; year of observation is 2000

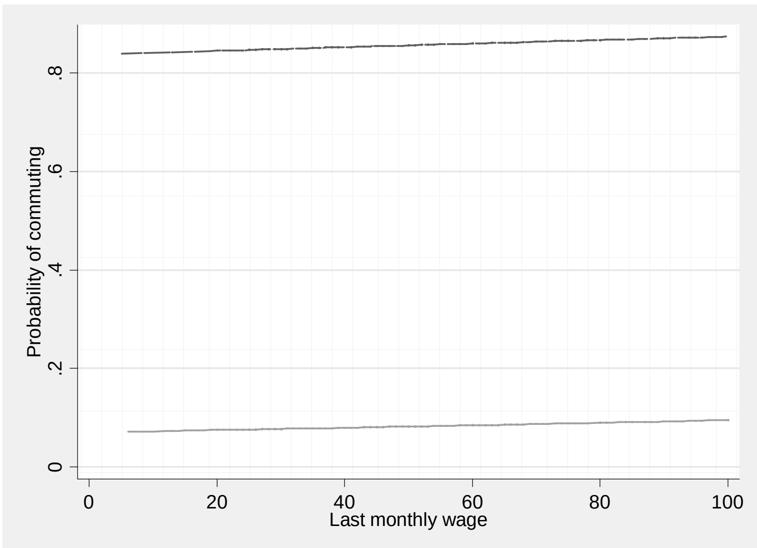
Figure 1
The predicted probability of commuting as a function of monthly gross wage by coverage of travel expenses

Legend: solid line: travel expenses not covered; dashed line: travel expenses covered

A) Men



B) Women



Notes

Predicted probabilities are calculated from the parameter estimates shown in *Table 6.* For both sexes, the curves depict the following situation: the employee is 40 years old and has A-level education, the local unemployment rate higher than 15 per cent, and the wage data is observed in year 2000.

Regional and Individual Determinants of the Willingness to Migrate in the Czech Republic

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Abstract

This paper adds to the literature attempting to explain low mobility in the accession countries by using the response to a question concerning the willingness to migrate in a large scale questionnaire on economic expectations and attitudes conducted in the Czech Republic. We find that variables measuring regional labour market conditions and amenities contribute little to explaining the willingness to migrate, but that personal and household characteristics such as income, residence in a family house and level of education are more important determinants.

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Introduction

Low internal migration rates in the Central and Eastern European candidate countries to the European Union have been the focus of a number of studies recently. In a comparative study Fidrmuc (2003) finds that overall internal mobility in candidate countries is low, has been falling over the last decade and is inefficient in reducing regional disparities. Ederveen and Bardsley (2003) find that migrants in the candidate countries are less responsive to regional wage and income disparities than in current EU member states and Drinkwater (2003a) using International Social Survey Programme data reports that of seven candidate countries considered (Bulgaria, Czech Republic, Hungary, Latvia, Poland, Slovakia and Slovenia) only Poland ranks in the upper half of a list of 20 countries' willingness to migrate. Cseres-Gergeley (2002), Hazans (2003), Kallai (2003) and Fidrmuc and Huber (2003) in a series of case studies on Hungary, the Baltics, Romania and the Czech Republic provide further evidence on low migration in candidate countries.

This paper adds to the literature attempting to explain the low mobility in these countries by using the response to a question concerning the willingness to migrate in a large scale questionnaire on economic expectations and attitudes conducted in the Czech Republic in April 1998. Our focus is on the personal and regional determinants of the willingness to migrate in the Czech Republic. In particular our aim is to identify regional or personal factors, which impede on willingness to migrate and to identify those groups of persons, who are most likely to be willing to migrate across regions. We find that variables measuring regional labour market conditions and amenities in general contribute little to explaining the willingness to migrate, but that personal and household characteristics such as income, residence in a family house and level of education are more important. We thus conclude that housing market imperfections and a low responsiveness to regional labour market disparities may be an important component to explaining low migration in accession countries. Furthermore, we find evidence that labour market conditions in neighbouring regions have a significant impact on the willingness to migrate. This may be evidence of commuting acting as a substitute for migration in a number of regions. Finally, we find substantial heterogeneity in the determinants of the willingness to migrate among subgroups. In contrast to males, higher education does not increase the willingness to migrate for females and for the less educated a longer duration of unemployment spells in the last two years reduces the willingness to migrate substantially, which is not the case for the overall smaple.

The paper is organised as follows: The next section presents the data used. Section 3 proceeds to present a model of the choice of answer to the question posed in our questionnaire on the willingness to migrate. We construct a matching model (see: Pissarides, 1990), and show that in such a model aside from individual factors influencing psychic and physical migration costs, regional factors such as wage disparities, labour market tightness and amenities will influence the willingness to migrate. Section 4 discusses our empirical approach and Section 5 presents results. Section 6 concludes.

Data

The data we use stem from the 11th Survey on Economic Expectations and Attitudes conducted in the Czech Republic in April 1998. In this survey a representative sample of 1075 individuals was interviewed on their households' financial and socio-economic position, employment experiences, their expectations of economic development for the next two years and their political attitudes and opinions concerning political reforms as well as the most important political debates in the Czech Republic at the time. Among the over 100 questions posed the one which is of interest to us is: "In case you would not have a job and you would have a possibility to get a job and a flat in another, distant municipality, would you be ready to move?". Respondents were given four options to answer. These were: definitely yes (encoded as 4 in our data), rather yes (3), rather not (2) and definitely not (1).

Table 1 presents the answers by selected personal characteristics and across regions. In total only 17.3% of the interviewed answered that they would definitely move if unemployed and offered work and residence in a distant region. A further 23.4% indicated that they would probably move. By contrast almost 31.0% of the interviewed stated that they would definitely not move and a further 28.4% stated that they would rather not move.

The data thus reconfirm the view that Czech citizen are in general unwilling to migrate for labour market reasons. This is also reconfirmed when focusing exclusively on the economically active (i.e. when excluding pensioners, housewives and students). Furthermore, descriptive statistics suggest that males and single persons are more willing to migrate, while less educated (in particular those having received elementary education) are substantially less willing to migrate.

Table 1: Distribution of Responses by selected personal and regional characteristics

	definitely not=1	rather not=2	rather yes=3	definitely yes=4	Nobs
All persons	30.98	28.37	23.35	17.30	1075
only economically active	27.84	31.11	23.98	17.08	855
male	27.41	29.30	24.57	18.71	506
female	34.43	27.47	22.16	15.93	523
Single	15.72	22.01	32.70	29.56	195
Married	33.60	30.53	20.67	15.20	750
Divorced	21.00	28.00	35.00	16.00	100
Widowed	53.03	19.70	13.64	13.64	66
Elementary	39.07	25.58	20.93	14.42	215
Vocational	29.21	30.94	25.50	14.36	404
Secondary	26.54	28.40	23.15	21.91	324
University, College	34.09	25.00	21.21	19.70	132
Prague	24.81	29.32	27.07	18.80	133
Central Bohemia	34.85	25.76	26.52	12.88	132
Southern Bohemia	31.43	35.71	21.43	11.43	70
Western Bohemia	23.96	26.04	35.42	14.58	96
Northern Bohemia	18.55	28.23	30.65	22.58	124
Eastern Bohemia	36.15	25.38	16.15	22.31	130
Southern Moravia	38.81	24.88	19.40	16.92	201
Northern Moravia	32.28	33.86	17.46	16.40	189

Furthermore, our data suggests substantial regional differences in the willingness to migrate. Persons, residing in Central Bohemia – the region bordering on Prague – and Southern Bohemia – a region located at the border to Germany and Austria – are least willing to migrate, while residents of Northern and Eastern Bohemia are much more willing to migrate. Although this regional variance could be explained by differences in composition of the workforce, coefficients of correlation suggest that at least some of this variance may be associated with differences in regional labour market conditions. The average willingness to migrate reported in table 1 is positively correlated with the regional unemployment rate and negatively with regional wages. After omitting the capital city of Prague from the sample (which may be a special case because of the substantial commuting into the area), the respective correlation coefficients are -0.43 and 0.52. These are insignificant, however, due to the small number of observations.

The Model

To model the choice of answer to the question under consideration we look at an economy consisting of a number of (I) regions, which are sufficiently distant from each other so that commuting is impossible and cast our discussion in the framework of a standard matching model of the labour market (see: Pissarides, 1990).[1] In each region (i) at time t unemployed persons receive unemployment benefits (b_t) and employed persons receive wages (w_{it}). Furthermore, employed persons face an exogenous probability of job loss of (s) in each period, while only the unemployed search for jobs with constant search intensity. Finally, the probability for an unemployed searcher to be matched to a job (p_{it}) in time period is determined by a matching function, which depends on the unemployment and vacancy rate in the region of residence i such that:

(1) $$p_{it} = \varphi(u_{it}, v_{it})$$

where u_{it} and v_{it} are the unemployment and vacancy rates, respectively.

Individuals derive utility from income and amenities (a_i) in region i. The expected utility of a risk neutral employed person living in region i is thus given by the returns of receiving wages (net of amenities in region i) and the expected value of future benefits of residing in region i, which depend on the chances of loosing or retaining the current job. The "asset value" of holding a job in region i (V_i) is thus given by:

(2) $$V_{it} = w_{it} + a_i + \frac{1}{1+\rho}[(1-s)V_{it+1} + sU_{it+1}]$$

with $(1+\rho)^{-1}$ a discount factor.

Similarly, an unemployed person receives unemployment benefits (net of amenities in region i) and the expected value of future benefits of residing in region i. Thus the asset value of being unemployed in region i (U_i) is given by:

(3) $$U_{it} = b_t + a_i + \delta[\varphi(u_{it+1}, v_{it+1})V_{it+1} + (1-\varphi(u_{it+1}, v_{it+1}))U_{it+1}]$$

In steady state U_{it} will be equal to U_{it+1} and V_{it} will equal V_{it+1} and both unemployment and vacancy rates will be independent of time.[2] Thus dropping time subscripts and solving (2) and (3) for V_i and U_i yields:

(4) $$\rho V_i = a_i + \frac{w_i(\rho + \varphi[u_i, v_i]) + sb}{\rho + s + \varphi[u_i, v_i]}$$

and

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[1] See Molho, 2001 and Jackman and Savvouri (1990) for applications of this model to spatial search and migration.

[2] See Pissarides, 1990 for a proof of existence in of a stable steady state in a model such as this.

(5) $$\rho U_i = a_i + \frac{b(\rho + s) + \varphi[u_i, v_i]w}{\rho + s + \varphi[u_i, v_i]}$$

Finally, if an individual (k) moves from region i to j we assume that it has to pay a cost of migration t_{ij}. These costs of migration vary across individuals and are determined by observable characteristics of the person (denoted by c^k), distance between the sending and receiving region (d_{ij}) and a random component ξ^k. In the question posed in the questionnaire respondents are put in front of the hypothetical situation of unemployment in their region of residence i. Thus given individual (k) is unemployed in region i and has a job offer in region j, as implied in our question, a risk neutral individual should prefer moving to staying in the region (be willing to migrate) if :

(6) $$V_j - U_i > \frac{t_{ij}(c^k, d_{ij}, \xi^k)}{\rho}$$

or

(7) $$a_j - a_i + \left[\frac{w_j(\rho + \varphi[u_j, v_j]) + sb}{\rho + s + \varphi[u_j, v_j]} - \frac{b(\rho + s) + w_i\varphi[u_i, v_i])}{\rho + s + \varphi[u_i, v_i]} \right] - \frac{t_{ij}(d_{ij}, c^k, \xi^k)}{\rho} > 0$$

Thus the willingness to migrate will be influenced by regional characteristics such as the wage level, unemployment and vacancy rates as well as measures for amenities in both receiving and sending regions and on the costs of migration, which in turn depend on the personal characteristics and distances between regions. Thus equation (7) gives a condition for when a risk neutral unemployed will be willing to migrate from region i to j migration from region i to j. Writing this condition more compactly we get::

(8) $$y^* = F(w_i, w_j, u_i, u_j, v_i, v_j, t_{ij}) > 0$$

To empirically implement this model, however, it has to be noted that the possible answers to the question in the questionnaire were definitely yes, rather yes, rather not and definitely not. Thus we cannot observe y* but rather only one of the four possible answers which are encoded 1 through 4 respectively.

We thus assume that all individuals for whom (8) was fulfilled answered either by selecting the answer definitely yes (i.e.4) or rather yes (i.e.3), and that all other people answered rather not or definitely not (i.e. 2 or 1). Furthermore, we assume that the two extreme answers occurred if either y* was highly positive (for definitely yes) or negative (for definitely not). Denoting as μ_1 and μ_2 the cut of levels between choosing category 4 and 3 and 3 and 2 respectively and normalising the cut off level for the choice between category 2 and 1 to zero, we can write the behavioural model underlying the choice of answer (y) by:

(9) y=4 if $y^* \geq \mu 2$
y=3 if $\mu 2 > y^* \geq \mu 1$
y=2 if $\mu 1 > y^* \geq 0$
y=1 if $0 \geq y^*$

Furthermore in the question under consideration no choice is given for the receiving region j. We thus assume that the individual considers an "average potential receiving region" as the appropriate receiving region. We calculate the characteristics of this "average potential receiving region" as the average of a particular indicator across all regions except the region of residence of the individual respondent and linearise (8). This yields:

(10) $y^* = \beta \ln X_i + \alpha \ln Y_i + \gamma \ln d + \lambda Z^k + \xi^k$

with Z^k_i a vector of individual characteristics of person k living in region i, X_i the regional characteristics which are measured as relative to the mean of the country (i.e. $X_i = \frac{1}{I-1} \frac{\hat{X}_i}{\sum_{j \neq i} \hat{X}_j}$

when $\hat{X}_i$ is the untransformed variable for the region under consideration), Y_i are neighbouring regions variables which are defined as $Y_i = \frac{1}{I-1} \frac{\frac{1}{K} \sum_{k \in S} \hat{X}_k}{\sum_{j \neq i} \hat{X}_j}$ (where S is the set of K regions bordering on region i), d is the average distance from all other regions given as $d = \frac{1}{I-1} \sum_{i \neq j} d_{ij}$ which can be interpreted as a measure of peripherality .

Thus equations (9) and (10) under the assumption that ξ^k as follows a logistic distribution define a standard ordered logit model of the choice of answer to the question analysed in this paper.[3]

Estimation Issues

There are a number of issues, which need clarification before estimating the model above. First of all, aside from both the sending and receiving regions' unemployment, wage and vacancy rate the

[3] Alternatively on could assume that errors in (10) are normally distributed which would lead to an ordered probit model. Since both ordered logit and probit models lead to similar results in most applications (see e.g. Greene, 2000) we focus only on logit estimates below.

model leaves open which further regional variables should be included to measure amenities and which individual variables should be included.

For the individual characteristics we follow the literature on the willingness to migrate in other countries and use gender, age, household structure (the number of economically active , number of children and number of pensioners in the household), highest completed education (elementary or less, vocational, secondary, university) and marital status (a dummy variable for married persons, divorced and widowed). These variables have proven to be of importance in a number of studies on the willingness to migrate (see: Ahn et al 1999, Yang, 2000 and Drinkwater, 2003), which all find that females and less educated persons are less willing to migrate as are married and older persons. Furthermore, we include variables to measure current household income and squared household income as well as an indicator concerning the type of residence of the household (family house as the base category, co-operative flat, rented flat, owner occupied flat and other) and a dummy which takes on the value of 1 if the interviewed owned a weekend house and zero else, because a number of authors have suggested that willingness to migrate may be lower among home owners (e.g. Hughes and Mc McCormick, 1987) or that persons with low income may be liquidity constrained and thus the relationship between willingness to migrate and income should be non-linear (e.g. Burda et al 1998). Furthermore, we include variables on the duration of unemployment experiences in the last two years because Jackman and Savouri, 1992 as well as Gross and Schoening, 1984 provide evidence that long term unemployed are less likely to migrate. Finally, in an extension of the baseline specification we also experiment with less conventional variables such as the preferences for a certain economic system (socialism, social market economy, market economy), and a subjective measure of poverty by considering a question in which respondents were asked, whether they consider themselves poor or not.

As measures of regional amenities we include measures of criminality in a region (murders per inhabitant), environmental quality (tons of emissions of hazardous wastes per square kilometre[4]) and variables measuring availability of public infrastructure (schools per 1000 inhabitants, hospital beds per 10000 inhabitants). Furthermore, as a measure of the distance of the region of

[4] These are measured as the sum of emissions of solids, SO2 and NOx in tons per square kilometre, disagregating the emissions by waste categories does not change results reported below. In particular emissions remain insignificant throughout

residence from the average receiving region we take the average distance between the capital city of the region of residence to all other regions' capital cities. Finally, we include the unemployment rate in a region of residence as well as the vacancy rate and the average wage level as indicators of the labour market situation.

Table 2 presents summary statistics for these variables. In general the sample seems to fit aggregate statistics rather well. For instance in our sample 51% of the interviewed are female. This accords with official statistics. There is, however, an under-representation of unemployed at the expense of an overrepresentation of both employed and inactive persons in our data. According to official statistics registered unemployment in the Czech Republic was at around 7.5% in 1998 but in our questionnaire only over 3% were unemployed. This may be explained by the usual differences which arise between interview based measures of unemployment and registered unemployment. Also in our data set 87% of the interviewed do not recollect having experienced any unemployment spell in the two years preceding the interview, but 3% claim to have had spells exceeding the length of one year. This accords well with studies on labour market flows in the Czech Republic (see: Storm and Terrell, 1997), which find low escape probabilities from unemployment and thus a relatively high long term unemployment rate.

Finally, 42% of the interviewed in our sample live in a family house and another 9% own their flat. While we are unable to check for the representativity of the sample with respect to house ownership, this does suggest that the share of owner occupied housing in the Czech Republic approaches EU levels. According to Eurostat the unweighted average share of owner occupied housing in the EU is at around 60% and lies below 50% in countries such as the Netherlands, Germany or Sweden.

Aside from reporting summary statistics for the overall sample table 2 also displays the average characteristics of the respondents who answered that they would either definitely move to another region and of those who stated that they would definitely not move. Comparing the characteristics of persons choosing these two extreme responses reconfirms that females, less educated and people who are not singles are in general less willing to migrate. Furthermore, those willing to migrate are in general younger, less likely to own a house and have a smaller household size than those unwilling to migrate. The average age of a respondent that stated to be willing to move is 37.6 years, while that of a respondent that stated that he or she would definitely not move is 49.3 years. Also those willing to move have fewer pensioners, active aged persons and children living in their household than those not willing to migrate and 57% of the people not willing to migrate own a family house, while of those willing to migrate only 33% own a family house. Finally, students (i.e. persons still receiving schooling) are less likely to be unwilling to move.

Table 2: Descriptive statistics

	Overall		Definitely Yes		Definitely No	
	Mean	Std.Dev	Mean	Std.Dev	Mean	Std.Dev
Age	42.87	14.49	37.61	13.90	49.35	14.75
		Gender				
Male[a)]	0.49	0.50	0.53	0.50	0.43	0.50
Female	0.51	0.50	0.47	0.50	0.57	0.50
		Education				
Elementary[a)]	0.20	0.40	0.17	0.37	0.25	0.43
Vocational	0.38	0.48	0.30	0.46	0.36	0.48
Secondary	0.30	0.46	0.39	0.49	0.26	0.44
University	0.12	0.33	0.14	0.35	0.14	0.34
Student	0.04	0.20	0.11	0.32	0.00	0.05
Ln(household income)	9.57	0.51	9.65	0.54	9.49	0.50
Ln (household income squared)	91.87	9.76	93.32	10.64	90.31	9.47
Married	0.70	0.46	0.61	0.49	0.76	0.43
Divorced	0.09	0.29	0.09	0.28	0.06	0.24
Widowed	0.06	0.24	0.05	0.22	0.11	0.31
No. of pensioners in Household	0.33	0.62	0.20	0.52	0.55	0.74
No. of children in Household	0.85	0.92	0.97	0.93	0.70	0.90
No. of active in household	1.63	0.88	1.84	0.95	1.42	0.98
		Type of Residence				
Family house[a)]	0.42	0.49	0.33	0.47	0.57	0.50
Co-operative Flat	0.15	0.36	0.17	0.38	0.11	0.32
Rented Flat	0.32	0.47	0.35	0.48	0.23	0.42
Own Flat	0.09	0.28	0.11	0.31	0.07	0.26
Other	0.03	0.16	0.04	0.19	0.02	0.13
Owns weekend house	1.23	0.42	1.21	0.41	1.24	0.43
		Unemployment duration in last two years				
less than two months[a)]	0.05	0.21	0.07	0.25	0.03	0.17
two months to one year	0.06	0.24	0.04	0.20	0.06	0.23
one year or more	0.03	0.16	0.02	0.15	0.02	0.15
not at all	0.87	0.34	0.87	0.34	0.89	0.32
		Preferred system				
Socialism[a)]	0.08	0.28	0.05	0.22	0.17	0.37
Social market Economy	0.62	0.49	0.54	0.50	0.61	0.49
Market Economy	0.29	0.46	0.41	0.49	0.23	0.42
		Poor family				
definitely yes[a)]	0.09	0.28	0.13	0.33	0.09	0.29
rather yes	0.29	0.45	0.28	0.45	0.30	0.46
rather no	0.44	0.50	0.36	0.48	0.42	0.50
definitely not	0.19	0.39	0.24	0.43	0.18	0.39
Ln(urater)	-0.15	0.53	-0.15	0.57	-0.13	0.48
Ln(region wage)	0.06	0.15	0.06	0.15	0.05	0.14
Ln(vacancy rate)	-0.08	0.54	-0.07	0.57	-0.10	0.54
Ln(murders per inhabitant)	0.98	0.74	1.05	0.75	0.92	0.68
Ln (emissions per sq. km)	-1.27	1.80	-1.20	1.80	-1.46	1.69
Ln(hospital bed per 10000 inh.)	1.11	1.13	1.13	1.20	1.19	1.06
Ln(schools per 10000 inh.)	1.26	0.54	1.26	0.58	1.32	0.50
Ln(average distance)	5.30	0.20	5.28	0.19	5.29	0.20
Ln(unemployment rate neighbours)	-0.04	0.29	-0.02	0.29	-0.06	0.28
Ln(wages neighbours)	0.01	0.05	0.01	0.05	0.01	0.05
Ln(vacancy rate neighbours)	0.07	0.43	0.08	0.40	0.06	0.43
Ln(murders per inh. neighbours)	1.05	0.49	1.07	0.50	1.01	0.48
Ln(emissions neighbours)	-1.05	1.36	-0.97	1.32	-1.14	1.34
Ln(hospital beds neighbours)	-0.06	0.19	-0.06	0.20	-0.08	0.20
Ln(schools neighbours)	1.14	0.75	1.17	0.78	1.20	0.73
Nobs	1071		332		184	

By contrast those willing and those unwilling to migrate seem to live in regions with relatively similar characteristics both in terms of amenities and labour market situation. The average regional unemployment rate among those willing to migrate is comparable to that of those not willing to migrate. The same applies to regional wage rates and to indicators of public infrastructure and average distance to other regions. Only indicators of environmental quality and criminality differ slightly between the two groups.

The second issue which needs discussion is the question of which regional breakdown should be used for estimation. Our data are coded at the level of NUTS 4 regions (called Okresy in Czech). These regions in average cover approximately 1000 square kilometres and have around 130.000 inhabitants. One of the assumptions in our model is that sending and receiving regions are far enough apart from each other to make commuting impossible. Clearly with regions of this size this assumption may be violated in a number of instances and Burda and Profit (1996) provide some indirect evidence of some commuting in the Czech Republic in early transition, in particular between regions which are contingent on each other. Good labour market conditions in neighbouring regions or better provision of public infrastructure may thus reduce the willingness to migrate. For this reason we include regional labour market conditions and ammenities in the average neighbouring region to deal with commuting. The summary statistics for these variables are presented in the bottom panel of table 2. Once more they suggest that aside from unemployment rates in the neighbouring districts, these variables do not discriminate well between persons willing and unwilling to move.

Finally, a technical estimation problem arises in our specification because we are merging information at two levels of aggregation (individual and regional). As pointed out for instance in Greene (2000) this will result in group wise heteroscedasticity of errors. This in turn will bias standard errors of the estimates. We thus choose to correct for this potential bias by estimating the model by maximum likelihood estimation under the assumption of such group wise heteroscedasticity.

Results

Table 3 shows the ordered probit results for the variables analysed. In column (1) we first focus on the role of individual characteristics in determining the willingness to migrate by replacing the regional variables by region fixed effects. Among the variables included in these regressions age, gender, education and house ownership are the most important determinants of the willingness to migrate. Older people are significantly less willing to migrate as are females.

Furthermore, people who have completed more than elementary education are significantly more willing to migrate. The effect of education on the willingness to migrate seems to be non-linear, however. Persons with completed secondary education are not significantly more willing to migrate than those with vocational training and persons with a university education do not differ significantly from those with only secondary or vocational education. This may be explained by the substantially better re-employment possibility for highly educated. If chances of re-employment are better for one group than for the other irrespective of residence the model presented in section 3 would ceteris paribus suggest a decline in the willingness to migrate. People still receiving education (students), however, are more willing to migrate than persons who have completed their education even after controlling for age differences.

Housing variables are another important influence on the willingness to migrate. According to our results owners of family houses have a significantly lower willingness to migrate than persons living in other residences. Other forms of residence (owner occupied apartments, rented houses or apartments, cooperative housing and others) do, however, not differ significantly from each other with respect to their inhabitants' willingness to migrate. This can be explained either by housing market inefficiencies which preclude the rapid sale of family houses without financial loss, or they could be due to self-selection of people less willing to migrate into family housing. Property in the form of houses, however, in general seems to be a deterrent to migration, because also owners of a weekend house tend to be less willing to migrate.

Household income and the number of economically active members in a household are further significant individual characteristics influencing the willingness to migrate. As found in a number of studies the connection between willingness to migrate and household income is non-linear. Persons with very low income are substantially less willing to migrate than persons with medium income and high income earners are also less willing to migrate.

The time spent in unemployment in the last two years, does not have a significant impact on the willingness to move. In particular persons, who were unemployed for more than a year in the two year period preceding the interview, have a willingness to migrate, which is only slightly smaller than that of persons, who were never unemployed. This accords with the results of Ahn et al

(1999), who also find that the discouragement effects of long term unemployment on search activities are not of particularly high relevance in explaining low willingness to migrate in Spain. Similarly, the number of children or pensioners residing in a household is an insignificant deterrent to the willingness to migrate. This is somewhat untypical in terms of the literature on actual migration decisions which tends to find that children are an impediment to migration and that married persons tend to be less willing to migrate than singles (e.g. Hunt, 2000), but accords with the findings of a number of studies on the willingness to migrate in other countries (see Ahn et al, 1999, and Drinkwater, 2003).

Table 3: Logit - Regression Results (dependent variable willingness to migrate)

	(1)		(2)		(3)	
	coeff	std. dev.	coeff	std. dev.	coeff	std. dev.
Age	-0.038***	0.007	-0.034***	0.007	-0.032***	0.007
Gender						
Male[a)]	0		0		0	
Female	-0.514***	0.124	-0.447***	0.101	-0.429***	0.105
Education						
Elementary[a)]	0		0		0	
Vocational	0.441**	0.177	0.386**	0.177	0.389**	0.180
Secondary	0.583***	0.189	0.525***	0.189	0.529***	0.196
University	0.535***	0.237	0.493**	0.234	0.470***	0.235
Student	1.311***	0.371	1.238***	0.303	1.283***	0.315
Ln(household income)	-6.325**	2.575	-5.809***	2.199	-5.821***	2.183
Ln (household income squared)	0.332**	0.133	0.308***	0.113	0.308***	0.110
Married	0.035	0.254	-0.005	0.259	0.047	0.273
Divorced	0.531*	0.300	0.438*	0.251	0.460*	0.265
Widowed	0.777*	0.413	0.411	0.469	0.419	0.491
No. of pensioners in Household	-0.043	0.141	-0.093	0.145	-0.069	0.148
No. of children in Household	-0.027	0.080	-0.043	0.075	-0.065	0.075
No. of active in household	0.309***	0.103	0.256**	0.100	0.256**	0.101
Type of Residence						
Family house[a)]	0		0		0	
Co-operative Flat	0.882***	0.191	0.796***	0.146	0.780***	0.143
Rented Flat	0.898***	0.161	0.804		0.804***	0.139
Own Flat	0.860***	0.240	0.725***	0.201	0.734***	0.202
Other	0.820**	0.385	0.795**	0.352	0.822**	0.346
Owns weekend house	-0.307*	0.159	-0.283**	0.144	-0.270*	0.140
Unemployment duration in last two years						
less than two months[a)]	0		0		0	
two months to one year	-0.603*	0.364	-0.511	0.342	-0.470	0.347
one year or more	-0.471	0.476	-0.407	0.507	-0.447	0.505
not at all	-0.226	0.293	-0.299	0.273	-0.239	0.273
Preferred system						
Socialism[a)]					0	
Social market Economy					0.773***	0.255
Market Economy					1.065***	0.283
Poor family						
definitely yes[a)]					0	
rather yes					-0.613***	0.235
rather no					-0.699***	0.243
definitely not					-0.771***	0.281
Ln(urater)			-0.427	0.335	-0.299	0.335
Ln(region wage)			-0.877	1.480	-0.268	1.489
Ln(vacancy rate)			0.158	0.146	0.173	0.146
Ln(murders per inhabitant)			0.101	0.119	0.059	0.119
Ln (emissions per sq. km)			0.010	0.054	0.056	0.817
Ln(hospital bed per 10000 inh.)			0.198	0.122	0.199*	0.121
Ln(schools per 10000 inh.)			-0.191	0.162	-0.258	0.162
Ln(average distance)			-0.805**	0.345	-0.868**	0.341
Ln(unemployment rate neighbours)			1.172***	0.369	1.230***	0.371
Ln(wages neighbours)			1.195	2.568	1.791	2.588
Ln(vacancy rate neighbours)			-0.042	0.264	-0.058	0.263
Ln(murders per inh. neighbours)			0.078	0.285	0.040	0.277
Ln(emissions neighbours)			-0.001	0.074	-0.019	0.071
Ln(hospital beds neighbours)			0.531	0.488	0.459	0.489
Ln(schools neighbours)			-0.305	0.190	-0.254	0.188
Diagnostics						
Pseudo R2	*0.12*		*0.08*		*0.09*	
CHI2			*486.01*		*512.86*	
			(37)		*(42)*	
Nobs	*1070*		*1070*		*1070*	
H0:Proportional log Odds			*0.50*		*0.12*	
Merge Categories						
2 and 3			*0.75*		*0.35*	
1 and 2			*0.00*		*0.02*	
3 and 4			*0.03*		*0.03*	

[a)] Reference Category, * (**) (***) signifies significance at the 10% (5%) 1% level respectively

In column (2) of table 3 we replace regional dummies with labour market indicators and measures for amenities in a region and in the neighbouring regions. This results in only minor changes to parameters relative to the estimates in column (1) and test statistics suggest that this leads only to a minor reduction in the goodness of fit in the specification. Among the measures for characteristics of a the region of residence, however, only the measure for average distance to other regions turns out to have a significant impact and some variables (unemployment rate, vacancy rate and the number of hospital beds) have an unexpected insignificant sign. Furthermore, among the variables for neighbouring regions only the unemployment rate is significant. This suggests that the overall impact of regional variables on the willingness to migrate is small in the Czech Republic and that among the regional characteristics which lead to a low willingness to migrate the most important is peripherality of the region under consideration. Also the significance of the neighbouring regions employment rate may be indication of the relevance of commuting as an alternative to migration to achieve regional mobility as also proposed in Burda and Profit (1996).

Overall the explicative power of our regressions, is small with the Pseudo R2 value lying between 11.8% and 8.5% in the regressions reported in columns (1) and (2). In column (3) we thus look for a number of further potential explanations for differences in the willingness to migrate. In particular we focus on individual characteristics such as the preference for the market system by entering dummy variables for persons, who stated that the preferred a market economy or a social market economy (with socialism as the base category) in answer to the question "Which type of economy do you prefer?". We also include a subjective measure of poverty from a question which read "Do you think you are a poor family?" as further explanatory variables. Among these attitudes towards the market economy play an important role. In general the more in favour of a market economy a person is the higher is its willingness to migrate. Also the subjective measure of poverty is of some importance. Persons who consider themselves members of a poor household are substantially more willing to migrate than people who do not.

Table 4: Marginal Effects of Equation (3)

	Choice =1		Choice=2		Choice=3		Choice=4	
Age	0.006***	0.001	0.001***	0.000	-0.004***	0.001	-0.004***	0.001
			Gender					
Male[a)]								
Female	0.087***	0.021	0.014**	0.005	-0.051***	0.013	-0.050***	0.012
			Education					
Elementary[a)]								
Vocational	-0.077**	0.035	-0.015*	0.008	0.046**	0.020	0.046**	0.023
Secondary	-0.102***	0.035	-0.024*	0.013	0.061***	0.021	0.066**	0.027
University	-0.088**	0.040	-0.026	0.018	0.052***	0.024	0.061*	0.035
Student	-0.191***	0.032	-0.119***	0.040	0.093***	0.011	0.217***	0.072
Ln(household income)	1.181***	0.439	0.185**	0.094	-0.697***	0.267	-0.669***	0.254
Ln (household income squared)	-0.063***	0.022	-0.010**	0.005	0.037***	0.014	0.035***	0.013
Married	-0.010	0.056	-0.001	0.008	0.006	0.033	0.005	0.031
Divorced	-0.085*	0.045	-0.026	0.021	0.051*	0.027	0.061	0.039
Widowed	-0.078	0.083	-0.024	0.040	0.047	0.049	0.055	0.073
No. of pensioners in Household	0.014	0.030	0.002	0.005	-0.008	0.018	-0.008	0.017
No. of children in Household	0.013	0.015	0.002	0.003	-0.008	0.009	-0.007	0.009
No. of active in household	-0.052**	0.020	-0.008**	0.004	0.031***	0.012	0.029***	0.012
			Type of Residence					
Family house[a)]								
Co-operative Flat	-0.138***	0.023	-0.052***	0.015	0.081***	0.013	0.109***	0.024
Rented Flat	-0.152***	0.023	-0.040***	0.013	0.089***	0.014	0.103***	0.022
Own Flat	-0.128***	0.030	-0.052**	0.021	0.075***	0.016	0.105***	0.036
Other	-0.137***	0.044	-0.065	0.041	0.078***	0.021	0.124***	0.065
Owns weekend house	0.055*	0.028	0.009*	0.005	-0.032*	0.017	-0.031**	0.016
			Unemployment duration in last two years					
less than two months[a)]								
two months to one year	0.103	0.081	0.000	0.011	-0.057	0.042	-0.046	0.029
one year or more	0.098	0.118	0.001	0.017	-0.054	0.060	-0.044	0.042
not at all	0.047	0.051	0.011	0.015	-0.028	0.031	-0.029	0.036
			Preferred system					
Socialism[a)]								
Social market Economy	-0.162***	0.054	-0.013*	0.007	0.092***	0.030	0.083***	0.027
Market Economy	-0.194***	0.044	-0.061**	0.025	0.111***	0.024	0.144***	0.045
			Poor family					
definitely yes[a)]								
rather yes	0.131**	0.052	0.007	0.006	-0.074***	0.027	-0.064***	0.023
rather no	0.144***	0.051	0.017**	0.007	-0.083***	0.027	-0.078***	0.028
definitely not	0.170**	0.066	-0.003	0.014	-0.092***	0.032	-0.075***	0.023
Ln(urater)	0.061	0.068	0.009	0.011	-0.036	0.040	-0.034	0.039
Ln(region wage)	0.054	0.302	0.009	0.048	-0.032	0.178	-0.031	0.172
Ln(vacancy rate)	-0.035	0.030	-0.005	0.005	0.021	0.018	0.020	0.016
Ln(murders per inhabitant)	-0.012	0.024	-0.002	0.004	0.007	0.014	0.007	0.014
Ln (emissions per sq. km)	0.003	0.011	0.000	0.002	-0.002	0.007	-0.001	0.006
Ln(hospital bed per 10000 inh.)	-0.040*	0.025	-0.006	0.004	0.024	0.015	0.023*	0.014
Ln(schools per 10000 inh.)	0.052	0.033	0.008	0.006	-0.031	0.020	-0.030	0.019
Ln(average distance)	0.176**	0.070	0.028**	0.013	-0.104**	0.042	-0.100**	0.039
Ln(unemployment rate neighbours)	-0.250***	0.076	-0.039**	0.016	0.147***	0.046	0.141***	0.042
Ln(wages neighbours)	-0.363	0.522	-0.057	0.088	0.214	0.312	0.206	0.297
Ln(vacancy rate neighbours)	0.012	0.053	0.002	0.008	-0.007	0.031	-0.007	0.030
Ln(murders per inh. neighbours)	-0.008	0.056	-0.001	0.009	0.005	0.033	0.005	0.032
Ln(emissions neighbours)	0.004	0.014	0.001	0.002	-0.002	0.009	-0.002	0.008
Ln(hospital beds neighbours)	-0.093	0.099	-0.015	0.016	0.055	0.059	0.053	0.057
Ln(schools neighbours)	0.052	0.038	0.008	0.006	-0.030	0.023	-0.029	0.021

Since coefficients in ordered logit estimates are difficult to interpret and the impact of a variable on the probability to answer in a particular category is hard to determine from regression coefficients, in table 4 we calculated marginal effects of the variables in the last estimate in table 3 on different response categories. For continuously measured variables these marginal effects have the interpretation of the percentage change in the probability of an otherwise average person to answer in one of the respective categories given a unit (one percent in the case of logarithmic variables) increase in the dependent variable. For dummy variables marginal effects measure the percent impact of the probability of answering in a particular category given a change of the dummy variable from zero to one for an individual with otherwise average characteristics.

These marginal effects reconfirm the view of a lower willingness to migrate among, older persons less educated and women. The coefficient on age for instance suggests that increasing the age of a person by 10 years increases the chance of answering that he would definitely not be willing to move by 6% while reducing the probability of being definitely willing to move by 4%. Women have an ceteris paribus 8.7% higher chance to respond that they are definitely not willing to move than men and the chances of a woman answering that she would definitely or rather move are 5% lower each than those of men.

Furthermore, the marginal effects suggest that owners of family houses are by between 12.8% to 15.2% more likely to answer that they would definitely not move than owners of other housing categories, while their likelihood to answer they would rather not move is between 4.0% to 6.5% higher. Similarly, people who are in favour of a market system are also more likely to answer that they either would be rather or definitely be willing to migrate, while regional variables aside from the average distance to other district capitals and the unemployment rate in the have no significant impact on the willingness to migrate in the aggregate. The marginal effects of these two regional variables, however, seem to be large. A 1% higher unemployment rate in neighbouring regions reduces the chances of being definitely unwilling to move by a quarter, while increasing the chances of being definitely willing to move by 14.0%. Increasing the average distance to other regions by 1% increases the chances of answering as definitely not willing to move by 17.6% and those of being rather not willing to move by 2.8%. The chances of answering as being definitely or rather willing to move increase by 10.0% and 10.4%, respectively.

Thus these results give further confirmation that housing market imperfections in particular in the market for family house owners and a low responsiveness of individuals to regional labour market conditions may explain part of the low migration in the Czech Republic. In particular the second fact could be associated with some commuting in the regions, since regional labour market conditions in neighbouring regions are at least of some importance in explaining low migration rates in candidate countries. Marginal effects, however, also suggest that a particularly low willingness to migrate in the Czech Republic is found in the remote regions of the country, thus suggesting that low migration may be a particular problem in the periphery.

We performed a number of tests to gauge the quality of fit and robustness of our results reported above.[5] In particular we conducted Hausmann tests to check for the appropriateness of the proportional log odds assumption underlying the logit model.[6] The test statistics suggest that the null of proportional log odds cannot be rejected for any of the models reported in table 3. Also we conducted a number of Hausmann tests, whether neighbouring categories of answers could be merged. As can be seen in the bottom panel of table 3 these tests suggest that neither answer category 1 (definitely not willing to move) and 2 (rather not willing to move) nor category 4 (definitely willing to move) and 3 (rather willing to move) could be merged. But the tests also suggest that the variables included in the model are not sufficient to discriminate well between the two intermediate answer categories. This suggests that the two intermediate categories cannot be well discriminated between, on the basis of the model presented above.[7]

We also included a number of further variables such as dummy variables for the immediate border regions to the east and the west, since some analysts suggest that the possibility of finding employment abroad may have been a deterrent for migration in the early phases of transition (see: Svejnar, 1999) and indicators for the settlement size in which a respondent lived, to check for robustness of our results. All these variables remained insignificant throughout and changes to parameters of existing variables were minimal (see table A4 in the Appendix).

[5]) The results of these are available from the authors upon request

[6]) This is equivalent to a test of the null-hypothesis that coefficients are not equal across answer categories. Proportional log odds are thus appropriate if the null of the test cannot be rejected.

[7] We performed a number of estimates in which these two categories were merged into one intermediate group and others where both intermediate categories were omitted. This, however, led to no further insights over the model reported here. Thus our preference was remaining with the original responses than moving to models with fewer answer categories.

Furthermore, we also experimented with including age squared and higher order terms for income, since Burda et al (1998) suggest that the relationship between income age and migration was not linear in the case of German East-West migration and propose including income cubed as a further explanatory variable in willingness to migrate regressions. These higher order terms remain insignificant throughout and in the case of age squared remove significance of the linear term (see table A4 in the Appendix). Finally we also experimented with including the current labour market status of the persons interviewed, from a fear that for instance the currently employed might show a different response behaviour than the currently unemployed, because of the hypothetical nature of the question posed. The dummy variables for inactive, unemployed relative to employed were insignificant, however.

Differences among Subgroups

A further issue which interested us was, whether different subgroups of the population react differently to certain influences on the willingness to migrate. We thus estimated the model separately for males, females and persons who completed only elementary education or less. Table 4 displays the coefficient estimates for each of these groups and suggest substantial variance with respect to the determinants of different subgroups' willingness to migrate.[8]

In particular for females higher educational attainment is not a significant determinant for the willingness to migrate, for males household size (in particular the number of economically active) is irrelevant and for the less educated – in contrast to the full sample – previous unemployment duration is an important determinant. Low educated persons with a longer unemployment duration in the last two years are less likely to migrate. This in turn suggests that discouragement effects play an important role in determining the low willingness to migrate among the less educated.

[8]) Marginal Effects are reported in tables A1 through A3 in the appendix

Table 5: Estimates for Subgroups

	male		female		low educ	
Age	-0.026**	0.009	-0.040***	0.010	-0.062***	0.016
Gender						
Male[a)]						
Female					0.356	0.309
Education						
Elementary[a)]						
Vocational	0.593**	0.250	0.223	0.267		
Secondary	0.966***	0.280	0.204	0.273		
University	0.963***	0.340	0.117	0.391		
Student	1.453***	0.509	1.082*	0.454	1.677**	0.763
Ln(household income)	-6.568	4.237	-6.407***	2.426	-6.758	6.485
Ln (household income squared)	0.342	0.216	0.341***	0.124	0.353	0.323
Married	-0.087	0.336	0.109	0.380	0.538	0.634
Divorced	0.113	0.408	0.675	0.432	0.248	0.651
Widowed	1.199*	0.679	0.303	0.590	0.042	0.941
No. of pensioners in Household	-0.017	0.173	-0.156	0.276	-0.026	0.441
No. of children in Household	0.008	0.111	-0.165	0.106	0.010	0.196
No. of active in household	0.214	0.143	0.257**	0.125	-0.201	0.240
Type of Residence						
Family house[a)]						
Co-operative Flat	0.992***	0.269	0.681***	0.224	1.175**	0.528
Rented Flat	0.931***	0.218	0.776***	0.226	1.420***	0.447
Own Flat	0.783**	0.314	0.833**	0.397	0.912**	0.512
Other	-0.018	0.573	1.354***	0.439	0.954	0.756
Owns weekend house	-0.336	0.219	-0.307	0.202	-1.071***	0.370
Unemployment duration in last two years						
less than two months[a)]						
two months to one year	-0.652	0.558	-0.367	0.453	-1.847***	0.857
one year or more	-0.776	0.717	-0.278	0.711	-2.363***	0.748
not at all	-0.057	0.453	-0.241	0.354	-1.243	0.728
Preferred system						
Socialism[a)]						
Social market Economy	0.838**	0.380	0.845**	0.381	0.897*	0.499
Market Economy	0.987**	0.410	1.260***	0.396	0.940*	0.525
Poor family						
definitely yes[a)]						
rather yes	-0.546	0.374	-0.698**	0.331	-0.355	0.490
rather no	-0.751**	0.379	-0.718**	0.346	-0.127	0.516
definitely not	-0.905**	0.428	-0.735**	0.357	0.167	0.598
Ln(urater)	0.042	0.430	-0.242	0.394	1.145	0.719
Ln(region wage)	0.042	2.016	0.756	1.803	-5.135	3.168
Ln(vacancy rate)	0.262	0.213	0.164	0.172	0.453	0.335
Ln(murders per inhabitant)	-0.304*	0.171	0.349***	0.134	0.565**	0.252
Ln (emissions per sq. km)	-0.058	0.082	-0.012	0.074	0.152	0.127
Ln(hospital bed per 10000 inh.)	-0.012	0.199	0.408***	0.153	-0.307	0.322
Ln(schools per 10000 inh.)	-0.724***	0.269	0.080	0.192	-0.336	0.409
Ln(average distance)	-1.301*	0.729	-0.649	0.433	-2.037*	1.073
Ln(unemployment rate neighbours)	1.215*	0.638	1.313***	0.431	-0.516	1.394
Ln(wages neighbours)	6.581*	3.912	-2.929	3.207	-3.694	6.895
Ln(vacancy rate neighbours)	-0.117	0.374	0.114	0.315	0.436	0.617
Ln(murders per inh. neighbours)	-0.219	0.377	0.272	0.356	0.896	0.652
Ln(emissions neighbours)	0.052	0.123	-0.097	0.079	0.059	0.247
Ln(hospital beds neighbours)	-0.012	0.668	0.761	0.596	1.873	1.216
Ln(schools neighbours)	-0.046	0.289	-0.414*	0.235	-0.041	0.457
Pseudo R2	*0.10*		*0.11*		*0.23*	
CHI2	*222.92*		*298.27*		*255.25*	
	(41)		*(41)*		*(39)*	
Nobs	*526*		*544*		*215*	
H0:Proportional log Odds	*0.09*		*0.27*		*0.00*	
Merge Categories						
2 and 3	*0.00*		*0.69*		*0.78*	
1 and 2	*0.59*		*0.32*		*0.89*	
3 and 4	*0.98*		*0.97*		*0.00*	

Also when focusing on subgroups of the population, we find that a number of further regional variables become significant. In particular higher criminality in a region increases the willingness to migrate among women and less educated, and neighbouring region unemployment is important only for females. Finally, marginal effects of variables significant in all three specification vary widely. For instance for males increasing age by 10 years increases the probability of being definitely unwilling to move by 5%, for females this marginal effect is 9% and for the less educated it is 14%. Similarly, for less educated ownership of an own family house has a much more negative effect on the willingness to migrate than for the other subgroups

Conclusion

This paper adds to the literature attempting to explain low mobility in the accession countries by using the response to a question concerning the willingness to migrate in a large scale questionnaire on economic expectations and attitudes conducted in the Czech Republic in April 1998. Our focus is on determining personal and regional determinants of the willingness to migrate. We find that variables measuring regional labour market conditions and amenities in general contribute little to explaining the willingness to migrate, but that personal and household characteristics such as income, residence in a family house and level of education are more important determinants. We thus conclude that housing market imperfections and a low responsiveness to regional labour market disparities may be an important component to explaining low migration. Furthermore, we find some evidence that labour market conditions in neighbouring regions have a significant impact on the willingness to migrate. This may be evidence of commuting acting as a substitute for migration in a number of regions.

Furthermore when moving to the determinants of the willingness to migrate for different subgroups, we find substantial heterogeneity. Education is an insignificant determinant of the willingness to migrate for females, and less educated – in contrast to the overall sample - experience a decline in their willingness to migrate with longer unemployment spells, which suggests important discouragement effects of unemployment. From a policy perspective this result suggests that aside from problems with inefficient housing markets, low migration in the accession countries may have to be targeted with different policies for different groups of persons. In terms of regional policy our results also suggest that peripheral regions may be a particular focus for migration related policies, since persons living in regions, which are more distant from the average receiving region have a significantly lower willingness to migrate than inhabitants of other regions. Clearly in these regions improving infrastructure may be among the most effective measures to increase migration.

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Figure 1: Correlation of Positive Answers with regional wage levels

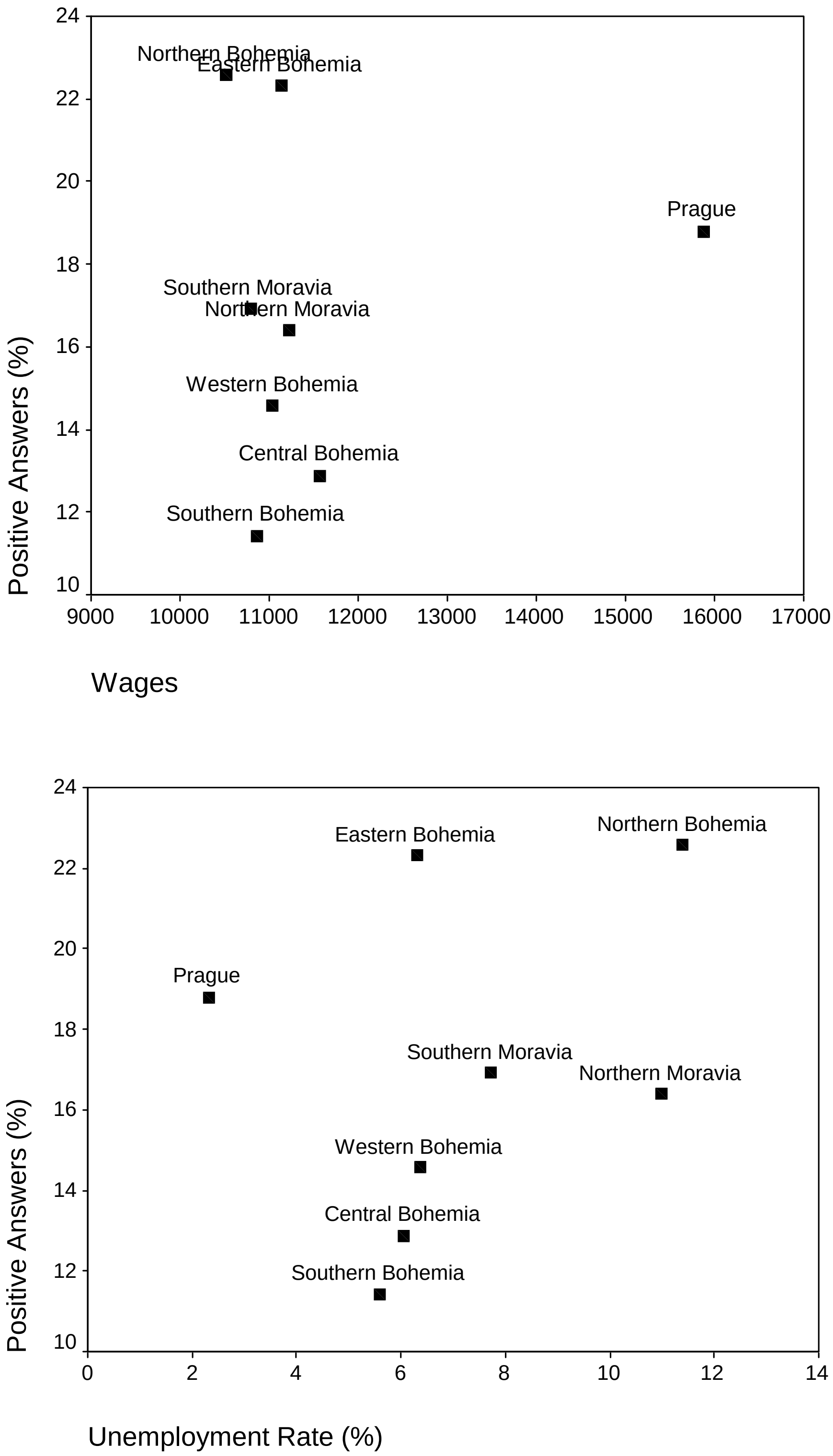

Table A1: Marginal Effects for Males see table 5

	Outcome = 1		Outcome=2		Outcome=3		Outcome=4	
	coeff.	std. err	coeff.	std. err	coeff.	std. err	coeff.	std. err
Age	0.005***	0.002	0.002**	0.001	-0.003**	0.001	-0.003***	0.001
Gender								
Male[a)]								
Female								
Education								
Elementary[a)]								
Vocational	-0.105***	0.040	-0.039**	0.018	0.067**	0.026	0.076**	0.032
Secondary	-0.154***	0.039	-0.083**	0.034	0.095***	0.025	0.141***	0.050
University	-0.141***	0.037	-0.095**	0.042	0.083***	0.019	0.153**	0.063
Student	-0.177***	0.045	-0.164**	0.063	0.072***	0.027	0.269**	0.120
Ln(household income)	1.188*	0.658	0.403*	0.233	-0.773*	0.418	-0.818*	0.467
Ln (household income squared)	-0.062*	0.033	-0.021*	0.012	0.040*	0.021	0.043*	0.024
Married	0.016	0.059	0.006	0.022	-0.010	0.038	-0.011	0.043
Divorced	-0.020	0.063	-0.008	0.027	0.013	0.041	0.015	0.048
Widowed	-0.155**	0.063	-0.133	0.103	0.076***	0.019	0.212	0.175
No. of pensioners in Household	0.003	0.035	0.001	0.012	-0.002	0.023	-0.002	0.024
No. of children in Household	-0.001	0.023	0.000	0.008	0.001	0.015	0.001	0.016
No. of active in household	-0.039	0.027	-0.013	0.010	0.025	0.018	0.027	0.019
Type of Residence								
Family house[a)]								
Co-operative Flat	-0.147***	0.031	-0.096***	0.034	0.087***	0.017	0.156***	0.055
Rented Flat	-0.153***	0.030	-0.074***	0.023	0.096***	0.020	0.131***	0.034
Own Flat	-0.118***	0.041	-0.075*	0.044	0.072***	0.022	0.121*	0.063
Other	0.003	0.098	0.001	0.032	-0.002	0.064	-0.002	0.066
Owns weekend house	0.061	0.047	0.021	0.016	-0.040	0.031	-0.042	0.031
Unemployment duration in last two years								
less than two months[a)]								
two months to one year	0.135	0.111	0.011	0.015	-0.080	0.062	-0.066*	0.039
one year or more	0.165	0.144	0.004	0.032	-0.095	0.073	-0.074*	0.042
not at all	0.010	0.078	0.004	0.030	-0.007	0.051	-0.007	0.057
Preferred system								
Socialism[a)]								
Social market Economy	-0.157**	0.077	-0.041**	0.018	0.098**	0.046	0.100**	0.049
Market Economy	-0.164**	0.069	-0.076*	0.042	0.102***	0.038	0.138**	0.074
Poor family								
definitely yes[a)]								
rather yes	0.105	0.085	0.023**	0.012	-0.066	0.050	-0.062	0.041
rather no	0.140*	0.084	0.039**	0.020	-0.088*	0.050	-0.090*	0.049
definitely not	0.183*	0.107	0.022**	0.012	-0.109*	0.056	-0.095**	0.041
Ln(urater)	-0.008	0.078	-0.003	0.026	0.005	0.051	0.005	0.054
Ln(region wage)	-0.008	0.356	-0.003	0.121	0.005	0.232	0.005	0.245
Ln(vacancy rate)	-0.047	0.039	-0.016	0.014	0.031	0.026	0.033	0.027
Ln(murders per inhabitant)	0.055	0.029	0.019*	0.011	-0.036*	0.019	-0.038*	0.020
Ln (emissions per sq. km)	0.011	0.014	0.004	0.005	-0.007	0.009	-0.007	0.010
Ln(hospital bed per 10000 inh.)	0.002	0.031	0.001	0.010	-0.001	0.020	-0.001	0.021
Ln(schools per 10000 inh.)	0.131***	0.042	0.044***	0.016	-0.085***	0.029	-0.090***	0.028
Ln(average distance)	0.235**	0.099	0.080**	0.035	-0.153**	0.066	-0.162**	0.067
Ln(unemployment rate neighbours)	-0.220**	0.095	-0.075	0.036	0.143***	0.061	0.151**	0.069
Ln(wages neighbours)	-1.191	0.604	-0.404	0.236	0.775*	0.409	0.820*	0.428
Ln(vacancy rate neighbours)	0.021	0.061	0.007	0.021	-0.014	0.040	-0.015	0.042
Ln(murders per inh. neighbours)	0.040	0.062	0.013	0.021	-0.026	0.040	-0.027	0.042
Ln(emissions neighbours)	-0.009	0.021	-0.003	0.007	0.006	0.014	0.006	0.014
Ln(hospital beds neighbours)	0.002	0.100	0.001	0.034	-0.001	0.065	-0.002	0.069
Ln(schools neighbours)	0.008	0.045	0.003	0.015	-0.005	0.030	-0.006	0.031

Table A2: Marginal Effects for Females see table 5

	Outcome = 1		Outcome=2		Outcome=3		Outcome=4	
	coeff.	std. err	coeff.	std. err	coeff.	std. err	coeff.	std. err
Age	0.009***	0.002	0.000	0.000	-0.005***	0.001	-0.004***	0.001
Gender								
Male[a)]								
Female								
Education								
Elementary[a)]								
Vocational	-0.048	0.056	-0.003	0.006	0.028	0.033	0.023	0.029
Secondary	-0.044	0.058	-0.002	0.005	0.025	0.034	0.021	0.029
University	-0.025	0.082	-0.002	0.008	0.015	0.048	0.012	0.042
Student	-0.187***	0.058	-0.076	0.053	0.108***	0.025	0.154*	0.089
Ln(household income)	1.392***	0.522	0.045	0.069	-0.803**	0.317	-0.634***	0.239
Ln (household income squared)	-0.074***	0.027	-0.002	0.004	0.043***	0.016	0.034***	0.012
Married	-0.024	0.084	0.000	0.001	0.014	0.048	0.011	0.037
Divorced	-0.131*	0.075	-0.030	0.032	0.079*	0.044	0.082	0.063
Widowed	-0.063	0.116	-0.008	0.025	0.037	0.071	0.033	0.070
No. of pensioners in Household	0.034	0.060	0.001	0.002	-0.020	0.035	-0.015	0.027
No. of children in Household	0.036	0.023	0.001	0.002	-0.021	0.013	-0.016	0.011
No. of active in household	-0.056**	0.027	-0.002	0.003	0.032**	0.016	0.025**	0.012
Type of Residence								
Family house[a)]								
Co-operative Flat	-0.134***	0.039	-0.027*	0.016	0.080***	0.026	0.081***	0.030
Rented Flat	-0.159***	0.043	-0.021	0.013	0.093***	0.027	0.086***	0.028
Own Flat	-0.155**	0.062	-0.045	0.037	0.093**	0.036	0.107*	0.063
Other	-0.216***	0.045	-0.110*	0.056	0.115***	0.019	0.211**	0.091
Owns weekend house	0.067	0.044	0.002	0.003	-0.038	0.025	-0.030	0.020
Unemployment duration in last two years								
less than two months[a)]								
two months to one year	0.084	0.108	-0.006	0.018	-0.046	0.055	-0.032	0.035
one year or more	0.063	0.167	-0.004	0.023	-0.035	0.087	-0.025	0.058
not at all	0.051	0.072	0.005	0.011	-0.030	0.043	-0.025	0.040
Preferred system								
Socialism[a)]								
Social market Economy	-0.190**	0.086	0.011	0.013	0.103**	0.044	0.076**	0.034
Market Economy	-0.237***	0.061	-0.060*	0.034	0.137***	0.034	0.160**	0.063
Poor family								
definitely yes[a)]								
rather yes	0.158**	0.077	-0.010	0.014	-0.086**	0.038	-0.062**	0.029
rather no	0.157**	0.077	0.001	0.008	-0.088**	0.041	-0.070**	0.035
definitely not	0.171**	0.087	-0.022	0.025	-0.089***	0.039	-0.060**	0.025
Ln(urater)	0.052	0.086	0.002	0.004	-0.030	0.049	-0.024	0.040
Ln(region wage)	-0.164	0.393	-0.005	0.012	0.095	0.226	0.075	0.176
Ln(vacancy rate)	-0.036	0.037	-0.001	0.002	0.021	0.022	0.016	0.017
Ln(murders per inhabitant)	-0.076***	0.029	-0.002	0.004	0.044**	0.017	0.035**	0.014
Ln (emissions per sq. km)	0.003	0.016	0.000	0.001	-0.001	0.009	-0.001	0.007
Ln(hospital bed per 10000 inh.)	-0.089***	0.034	-0.003	0.004	0.051**	0.020	0.040***	0.015
Ln(schools per 10000 inh.)	-0.017	0.042	-0.001	0.001	0.010	0.024	0.008	0.019
Ln(average distance)	0.141	0.094	0.005	0.007	-0.081	0.055	-0.064	0.043
Ln(unemployment rate neighbours)	-0.285***	0.094	-0.009	0.014	0.164***	0.055	0.130***	0.045
Ln(wages neighbours)	0.636	0.701	0.020	0.032	-0.367	0.401	-0.290	0.317
Ln(vacancy rate neighbours)	-0.025	0.068	-0.001	0.003	0.014	0.040	0.011	0.031
Ln(murders per inh. neighbours)	-0.059	0.077	-0.002	0.004	0.034	0.044	0.027	0.036
Ln(emissions neighbours)	0.021	0.017	0.001	0.001	-0.012	0.010	-0.010	0.008
Ln(hospital beds neighbours)	-0.165	0.130	-0.005	0.009	0.095	0.075	0.075	0.060
Ln(schools neighbours)	0.090*	0.052	0.003	0.004	-0.052*	0.031	-0.041*	0.022

Table A3: Maginal Effects for Low Education see table 5

	Outcome = 1		Outcome=2		Outcome=3		Outcome=4	
	coeff.	std. err	coeff.	std. err	coeff.	std. err	coeff.	std. err
Age	0.014***	0.004	-0.001	0.002	-0.009***	0.002	-0.004***	0.001
Gender								
Male[a)]								
Female	-0.079	0.067	0.009	0.010	0.050	0.044	0.020	0.019
Education								
Elementary[a)]								
Vocational								
Secondary								
University								
Student	-0.277***	0.085	-0.110	0.098	0.216***	0.068	0.171	0.120
Ln(household income)	1.501	1.445	-0.159	0.249	-0.956	0.910	-0.387	0.367
Ln (household income squared)	-0.078	0.072	0.008	0.013	0.050**	0.045	0.020	0.018
Married	-0.120	0.145	0.014	0.023	0.075	0.085	0.030	0.037
Divorced	-0.053	0.134	0.002	0.009	0.036	0.095	0.016	0.046
Widowed	-0.009	0.207	0.001	0.017	0.006	0.134	0.002	0.055
No. of pensioners in Household	0.006	0.098	-0.001	0.010	-0.004	0.062	-0.001	0.025
No. of children in Household	-0.002	0.044	0.000	0.005	0.001	0.028	0.001	0.011
No. of active in household	0.045	0.053	-0.005	0.007	-0.028	0.035	-0.011	0.014
Type of Residence								
Family house[a)]								
Co-operative Flat	-0.216***	0.074	-0.051	0.059	0.167**	0.071	0.100*	0.060
Rented Flat	-0.291***	0.079	-0.004	0.035	0.197***	0.065	0.098***	0.037
Own Flat	-0.174**	0.078	-0.032	0.050	0.133*	0.072	0.072	0.054
Other	-0.174	0.106	-0.045	0.084	0.138	0.102	0.081	0.087
Owns weekend house	0.238***	0.082	-0.025	0.028	-0.152***	0.057	-0.061**	0.024
Unemployment duration in last two years								
less than two months[a)]								
two months to one year	0.431**	0.170	-0.192*	0.113	-0.181***	0.058	-0.058***	0.017
one year or more	0.518***	0.115	-0.258***	0.085	-0.198***	0.043	-0.061***	0.016
not at all	0.231**	0.107	0.048	0.072	-0.176*	0.097	-0.103	0.081
Preferred system								
Socialism[a)]								
Social market Economy	-0.204*	0.118	0.037	0.036	0.120*	0.066	0.047*	0.024
Market Economy	-0.184**	0.090	-0.023	0.040	0.136*	0.078	0.071	0.048
Poor family								
definitely yes[a)]								
rather yes	0.080	0.112	-0.010	0.020	-0.049	0.067	-0.020	0.027
rather no	0.028	0.116	-0.003	0.016	-0.018	0.072	-0.007	0.028
definitely not	-0.036	0.127	0.002	0.004	0.024	0.088	0.010	0.038
Ln(urater)	-0.254	0.159	0.027	0.033	0.162	0.104	0.066	0.042
Ln(region wage)	1.141	0.697	-0.121	0.136	-0.726	0.463	-0.294	0.204
Ln(vacancy rate)	-0.101	0.073	0.011	0.013	0.064	0.049	0.026	0.020
Ln(murders per inhabitant)	-0.125**	0.056	0.013	0.015	0.080**	0.037	0.032*	0.017
Ln (emissions per sq. km)	-0.034	0.028	0.004	0.005	0.022	0.018	0.009	0.008
Ln(hospital bed per 10000 inh.)	0.068	0.070	-0.007	0.009	-0.043	0.046	-0.018	0.020
Ln(schools per 10000 inh.)	0.075	0.091	-0.008	0.013	-0.048	0.058	-0.019	0.023
Ln(average distance)	0.453*	0.238	-0.048	0.055	-0.288*	0.158	-0.117*	0.068
Ln(unemployment rate neighbours)	0.115	0.308	-0.012	0.032	-0.073	0.198	-0.029	0.081
Ln(wages neighbours)	0.821	1.524	-0.087	0.163	-0.523	0.983	-0.211	0.410
Ln(vacancy rate neighbours)	-0.097	0.137	0.010	0.019	0.062	0.088	0.025	0.035
Ln(murders per inh. neighbours)	-0.199	0.143	0.021	0.024	0.127	0.095	0.051	0.041
Ln(emissions neighbours)	-0.013	0.055	0.001	0.006	0.008	0.035	0.003	0.014
Ln(hospital beds neighbours)	-0.416	0.269	0.044	0.054	0.265	0.174	0.107	0.076
Ln(schools neighbours)	0.009	0.101	-0.001	0.011	-0.006	0.065	-0.002	0.026

Table A4: Alternative Specifications

	(1)		(2)		(3)	
	coeff.	std. err	coeff.	std. err	coeff.	std. err
Age	-0.032***	0.007	0.000	0.031	-0.031***	0.007
			0.000	0.000		
Gender						
Male[a)]						
Female	-0.430***	0.106	-0.429***	0.104	-0.429***	0.107
Education						
Elementary[a)]						
Vocational	0.392**	0.180	0.388**	0.185	0.391**	0.180
Secondary	0.529***	0.197	0.536***	0.199	0.508**	0.203
University	0.480**	0.235	0.471**	0.239	0.445*	0.237
Student	1.282***	0.314	1.412***	0.339	1.305***	0.322
Ln(household income)	-5.847***	2.164	20.268	30.775	-5.768***	2.077
Ln (household income squared)	0.310***	0.109	-2.436	3.189	0.306***	0.105
			0.096	0.110		
Married	0.054	0.275	0.010	0.265	0.045	0.273
Divorced	0.460*	0.264	0.369	0.259	0.427	0.263
Widowed	0.426	0.494	0.424	0.502	0.395	0.486
No. of pensioners in Household	-0.064	0.148	-0.052	0.158	-0.095	0.153
No. of children in Household	-0.064	0.075	-0.073	0.078	-0.063	0.076
No. of active in household	0.256**	0.101	0.253***	0.101	0.244**	0.102
Type of Residence						
Family house[a)]						
Co-operative Flat	0.787***	0.145	0.767***	0.144	0.777***	0.160
Rented Flat	0.819***	0.143	0.793***	0.139	0.811***	0.157
Own Flat	0.751***	0.201	0.732***	0.202	0.702***	0.206
Other	0.841**	0.346	0.803**	0.344	0.843***	0.343
Owns weekend house	-0.273*	0.141	-0.261*	0.142	-0.248*	0.144
Unemployment duration in last two years						
less than two months[a)]						
two months to one year	-0.478	0.346	-0.455	0.350	-0.384	0.357
one year or more	-0.462	0.506	-0.454	0.504	-0.396	0.490
not at all	-0.236	0.273	-0.232	0.274	-0.167	0.288
Preferred system						
Socialism[a)]						
Social market Economy	0.760***	0.256	0.746***	0.259	0.789***	0.253
Market Economy	1.047***	0.282	1.036***	0.285	1.093***	0.282
Poor family						
definitely yes[a)]						
rather yes	-0.613***	0.235	-0.641***	0.240	-0.599**	0.233
rather no	-0.702***	0.243	-0.702***	0.245	-0.675***	0.246
definitely not	-0.762***	0.282	-0.768***	0.283	-0.771***	0.283
Ln(urater)	-0.332	0.332	-0.290	0.336	-0.362	0.329
Ln(region wage)	-0.154	1.438	-0.241	1.509	0.013	1.579
Ln(vacancy rate)	0.197	0.138	0.171	0.146	0.087	0.138
Ln(murders per inhabitant)	0.048	0.111	0.065	0.120	0.083	0.117
Ln (emissions per sq. km)	-0.016	0.056	-0.014	0.057	-0.005	0.055
Ln(hospital bed per 10000 inh.)	0.226*	0.119	0.193	0.122	0.156	0.124
Ln(schools per 10000 inh.)	-0.286*	0.163	-0.249	0.163	-0.266*	0.161
Ln(average distance)	-0.634*	0.342	-0.886**	0.343	-0.997***	0.347
Ln(unemployment rate neighbours)	1.280***	0.375	1.221	0.379	1.213***	0.345
Ln(wages neighbours)	1.715	2.616	1.647	2.670	1.291	2.654
Ln(vacancy rate neighbours)	-0.034	0.253	-0.056	0.266	-0.204	0.247
Ln(murders per inh. neighbours)	-0.007	0.276	0.042	0.276	0.038	0.271
Ln(emissions neighbours)	-0.006	0.072	-0.018	0.072	-0.030	0.076
Ln(hospital beds neighbours)	0.662	0.513	0.488	0.492	0.499	0.507
Ln(schools neighbours)	-0.325*	0.195	-0.259	0.186	-0.235	0.193
neigheast	-0.151	0.160				
neighwest	-0.225	0.187				
					0.436	0.274
					0.490*	0.262
					0.260	0.241
					0.011	0.336
Pseudo R2	*0.09*		*0.09*		*0.09*	
CHI2	*525.81*		*539.40*		*583.45*	

	(1)	(2)	(3)
	(44)	*(44)*	*(46)*
Nobs	*1070*	*1070*	*1070*
H0:Proportional log Odds	*0.12*	*0.30*	*0.13*
Merge Categories			
2 and 3	*0.68*	*0.06*	*0.72*
1 and 2	*0.00*	*0.00*	*0.09*
3 and 4	*0.04*	*0.00*	*0.89*

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Looking for the Workforce: the Elderly, Discouraged Workers, Minorities, and Students in the Baltic Labour Markets

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Abstract

This paper looks at the evolution of the labour markets in Estonia, Latvia, and Lithuania since the beginning of transition (in some respects since 1996/1998) until 2003, with a particular focus on labour force participation. How did labour supply in the Baltic countries respond to changes in to minimum wages, unemployment benefits and retirement regulation? Do the marked differences in labour market policies between the countries result in different patterns of participation? What are the obstacles to and driving forces of participation?

We find that relative contribution of participation and demographic trends to the dynamics of the labour force varied substantially both over the years and across the three countries. Participation, in turn, has been shaped by sometimes complicated interaction between educational choices, retirement, policy changes, and external shocks. Resulting differences in trends and patterns are quite substantial, indicating that there is a room for increasing participation in each of the countries.

Recent rates of transition from unemployment to employment and to inactivity are similar to those found in EU-15.

Panel data analysis of determinants of participation and discouragement suggests that increasing after-tax real minimum wage has significant positive effect on participation and reduces discouragement in Lithuania. In Estonia, by contrast, positive effect of minimum wage on participation is found only for teenagers of both genders and for young males.

Ethnic minorities, especially females, in all three Baltic countries are less likely to be in the labour force, other things equal.

Key words: Labour supply; discouraged workers; labour market flows; minimum wages; ethnic minorities.

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Introduction

Three reasons have motivated this paper. First, labour force mobilisation is one of the ways to rise employment-population ratio, which in the Baltic countries is currently well below EU-15 average, let alone the Lisbon targets. But in the Baltic countries the importance of rising participation is reinforced by demographic factors. Hence it is urgent to understand the patterns of labour supply and to identify obstacles and possible incentives for specific groups. Second, the three neighbour countries have adopted different labour market policies with respect to minimum wage, unemployment benefits, and old-age pension, three issues clearly related to labour supply. How are these differences reflected in labour market outcomes is a policy relevant question. This introduction provides a more detailed discussion of the two above mentioned reasons behind the paper. The third reason is related to sizable ethnic minorities (mostly Russian speaking) which exist in the Baltic countries. Previous studies (see Kroncke and Smith (2000), Chase (2001), OECD (2003a-2003b), Hazans (2004b) have found that labour market outcomes (unemployment risk and earnings) are less favourable for ethnic minorities than for majority population. We shall test whether recent data support this conclusion with respect to labour force participation.

Effective policy making in the Baltic countries even more than in other countries of Central and Eastern Europe (CEE) is confounded by demographic trends. Figure 1, which displays combination of natural increase and net migration between 1989 and 2002, documents that Estonia and Latvia are the only countries in the region which experienced both negative natural increase and significant loss of population due to net migration. In Lithuania demographic boom of 1980s went on in 1990-1992, resulting in total positive natural increase over the period; however, in 1993 fertility slowed down, and since 1994 natural increase is negative, while net migration has been negative during the whole period. Overall, by the beginning of 2004, population of Estonia, Latvia and Lithuania went down by 13.7, 13.0 and 6.2 percent respectively, compared to 1989. In 2003, of European countries only Bulgaria experienced larger depopulation than the Baltic countries (Eurostat, 2004).

Given double-digit unemployment, in the short term labour shortage would not be a problem from the natural demographic perspective alone, because of comparatively large youth cohorts about to enter the labour force over the coming decade – before the effects of population ageing begin to have stronger influence (OECD 2003a). However, emigration, which has slowed down in 2001-2002, is likely to increase substantially in the years to come when restrictions on labour mobility between new and old EU member states will be gradually removed. While there is still a good deal of uncertainty about the size of emigration, the fact that Baltic labour force is among the most educated in the EU-25 (see Table 1), combined with still low (especially in Latvia and Lithuania) average earnings makes to think that outflow of labour will not be negligible.

Preliminary research (Hazans, 2003a; 2003b) suggests that (i) Baltic population seems to be relatively mobile in comparison with other European nations; (ii) on the eve of accession significant proportions of skilled non-manual, clerical and service workers, and students (the survey was limited to Internet users) seriously considered the possibility of moving permanently or temporarily to one of the EU countries if this were possible. Available bits of post-accession evidence confirm these expectations and suggest that also many manual workers are looking west. In Latvia recently launched bus line to Ireland (one of the few restriction-free EU-15 countries) is booming, and flights to Ireland are in big demand, too. Table 2 presents official UK data on registered immigrants from the new EU members during the first 6 post-accession months, adjusted to countries' population figures. Lithuania and Latvia top the list very convincingly; Estonia, though slightly below Poland and Slovakia, still features a rate two times higher than Czech Republic and four times higher than Hungary.

According to UN/ILO projections, demographic limitations on labour supply are set to become gradually more critical in the years after 2015, and by 2040 the ratio of persons aged 65 or more to population aged 20 to 64 is going to almost double compared to the year 2000 level; in reality ageing might be even more pronounced because the projections for the post-accession emigration, which is likely to be "young". The OECD (2003a) report warns Baltic countries that "insofar as a possibly emerging scarcity of labour in the future would be unlikely to be offset by a steep rise in immigration or fertility, it will be all the more important to enhance the existing human capital and to ensure that it is productively employed".

This paper aims at identifying important patterns of labour force participation (including the discouraged worker effect) in the three Baltic countries, as well as relating the findings to the marked differences in unemployment benefit and minimum wage policies.

Figure 2 displays evolution of proportion of unemployment benefits (UB) recipients among registered unemployed, along with evolution of average UB – average wage ratio in the Baltic countries[2]. Of the three countries Latvia has the most generous UB system, which

[2] In Estonia (until 2002) and in Lithuania UB were not taxed, so the ratio of UB to average net wage is used. For Latvia, where UB are taxed, Figure 2 shows average UB – average gross wage ratio (the ratio of after-tax UB to net wage would be almost identical).

covered about 30 percent of registered unemployed prior to 1999 and more than 40 percent since then, with average UB between 25 and 30 percent of average wage in most years. In Lithuania the relative level of average UB has been roughly same as in Latvia until 2001 and somewhat higher since then, reaching 36 percent in 2003 due to special treatment of the elderly; however, the coverage in Lithuania since 1997 has been much lower than in Latvia and falling every year, with just 11 percent covered in 2003. Another important difference is that in Latvia UB are earnings related, while in Lithuania they depend only on number of years of contribution. In Estonia, before 2003 UB have been paid at a flat rate and in most years covered 49 to 60 percent of registered unemployed. Initially, in 1992, UB amounted to 31 percent of average net wage but this ratio felt sharply to less than 10 percent by 1995 and then varied between 6.4 and 11.4 percent until 2002. In 2003 new unemployment insurance system has started to pay benefits, raising total coverage to 76 percent, and overall average UB - wage ratio to 16 percent. More details on UB in the Baltic countries are found in Table A1.

Both levels and dynamics of minimum wage also have been very different in the three countries (see Table 3). The ratio of minimum to gross average wage in Estonia dropped from 36 to 19 percent between 1992 and 1995; since then it has been gradually increasing and reached 32 percent by 2003, with nominal minimum wage changing once a year. In Latvia the same ratio has increased from 27 to 36 percent between 1992 and 1996; since then it has been fluctuating between 31 and 36 percent, with nominal minimum wage changing typically every second year (recently adopted new policy envisages annual adjustments in future). In Lithuania a major change took place between 1994 and 1997, when the minimum wage - average wage ratio has increased from 17 to 48 percent; since then it has declined to 41 percent, yet it is well above the ratio found in Estonia and Latvia; the last change in the nominal level of minimum wage took place in 1998, while in 2002-2003 non-taxable minimum has been raised instead. On top of these differences, there is substantial variation of minimum wage-average wage ratio across the regions in each country, due to inter-regional wage differentials (see Hazans 2003a for details).

The rest of the paper is organised as follows. Section 2 briefly surveys the literature and relates this paper to previous studies. Section 3 provides a comparative analysis of major trends in labour force participation in the three countries, focusing on annual changes in population, employment, unemployment, and inactivity of population aged 15-64, as well as of those aged 65-74; the latter group is of course of a special interest as a potential reserve for labour force mobilisation. Section 4 amends this analysis by looking at flows between employment, unemployment, and inactivity. Section 5 revises age and gender related trends and patterns of labour force participation. Sections 6 and 7 provide an econometric analysis of determinants of labour force participation and discouragement, using panel data from recent Labour Force Surveys. Section 8 concludes.

Survey of the literature

Labour supply in transition countries has been subject of extensive research (see Svejnar (1999) and Huber et al (2002) for detailed surveys). Simple decomposition of changes in employment rates has led to conclusion that in some countries, like Hungary, Czech Republic and Bulgaria, reduced participation has been a major factor in declining employment in 1990-1996, while it played a minor role in other countries (Boeri, Burda, Kollo, 1998). Studies of flows between employment, unemployment and inactivity found, among other things, that flows into inactivity have represented a substantial part of the adjustment mechanism, while probabilities of transition from inactivity are lower than in matured market economies (Storm and Terrell, 2000; Boeri, 2001). According to Boeri (2001), Boeri and Terrell (2002) disincentive effects of non-employment benefits play important role in individual labour supply decisions and, accordingly, in shaping the labour market flows; Boeri (2001) has suggested a model which incorporates these effects.

Previous research of labour supply in the Baltic countries has been largely limited to studies of flows between employment, unemployment and inactivity in papers and reports whose main focus was other than labour supply. Haltiwanger and Vodopivec (2002) analyse annual flows for Estonia 1989-1995; OECD (2003a, 2003b), relying on Hazans, Earle and Eamets (2002), inspects ten years flows between 1990 and 2000, as well as annual flows for Estonia, Latvia (1997-2000) and Lithuania (1999-2000); these annual flows are further discussed by Eamets (2004) in the context of adjustment to macroeconomic shocks. Rutkowski (2003) and Hazans (2004b) analyse annual flows in Lithuania (2000-2001) and Latvia (2000-2002) respectively. Descriptive analysis of labour force participation in the Baltic countries is found in OECD (2003a, for 1997-2000), Rutkowski (2003, for Lithuania, 1997-2001), Hazans (2004a, Latvia, 1996-2002). Econometric analysis of determinants of labour force participation in Latvia is provided by Chase (2001) and Hazans (2004b). Eamets (2004) looks at simultaneous annual changes in employment, unemployment and inactivity in the late 1990s and finds some evidence for discouraged worker effect in Latvia and Lithuania but not in Estonia – a finding which is modified in this paper via more detailed analysis. This paper will take a unified view on the existing evidence, adding also more recent Lithuanian flows (2002-2003).

As far as minimum wages are concerned, recent studies by Hinnosaar and Room (2003) and Kertesi and Kollo (2003) have found disemployment effect of increasing minimum wage in Estonia and Hungary, but this seems to be a demand side effect. Kollo (2001) have found no conclusive evidence on minimum wage effect on labour force participation.

Accounting the reallocation of labour

We start with looking at the major labour market trends in each of the three Baltic countries during the period from 1989 to 2003. Evolution of population, labour force, employment and real GDP is presented in Figure 3. Initial output decline, from nearly 50 percent in Latvia to 35 percent in Estonia, was substantially deeper than elsewhere in Central and Eastern Europe. While GDP decline has been reversed in 1995, labour force continued to fall faster than population until 1999 in Estonia; in Latvia and Lithuania this pattern prevailed until 2000 and 2001 respectively. Two or three years earlier, however, negative trend in employment has been either temporarily reversed (in Latvia, 1997 and Lithuania, 1998) or muted (in Estonia, 1997).

This suggests a natural breakdown of the whole transition period into three episodes:

(i) From the beginning of the transition until 1996 or 1997, when both labour force and employment were declining (this was also a period of growing unemployment);

(ii) A three or four year period from the initial recovery of employment in 1997 or 1998 until the end of labour force contraction period. Except for the first year in Latvia and Lithuania, this was also a period when employment and labour force were declining, although much slower than in 1992-1995. Unemployment trends were mixed (see below). The second part of this episode includes the period when the three Baltic economies were heavily affected by the Russian financial crisis of 1998. Negative GDP growth was observed, however, only in 1999 in Estonia and Lithuania.

(iii) A period of recovery of employment in 2001-2003 (for Lithuania, 2002-2003), with generally declining unemployment but mixed trends in participation.

Table 4 decomposes changes in labour force during each of the three sub-periods into contributions from trends in demographics and participation rates to labour force. Likewise, changes in employed population are tracked down to changes in demographics, participation rates, and unemployment rates. These results follow from the identities

$$LF = \frac{LF}{POP_{15-64}} \frac{POP_{15-64}}{POP} POP \, , \quad E = (1-u)LF \, , \qquad (1)$$

where LF is number of members of the labour force, POP and $POP_{15\text{-}64}$ – total population and population aged 15 to 64, E – number of employed persons, u – unemployment rate. Note that 97 to 99 percent of the labour force comes from the 15-64 age group, hence proportion of this group in population is an important determinant of labour supply.

Findings from Table 4 can be summarised as follows. Contraction of the labour force between 1989 and 1996/7 was almost 20 percent in Estonia and Latvia; demographic trends and declining participation contributed almost equally to this contraction. In Lithuania, by contrast, labour force declined in the same period by less than 12 percent, of which 8 percent were due to change in participation. Declining labour force and increasing unemployment rate contributed almost equally to fall in the number of employed persons in Latvia and Lithuania, while in Estonia contracting labour force was responsible for two thirds of the total change in employment. In this respect the Baltic countries are similar to Hungary, Bulgaria, and Czech R. (see Boeri, Burda, and Kollo, 1998), but demographic trends were much more important in the Baltic.

During next three or four years (encompassing the Russian crisis), labour force has declined further by 3 percent in Estonia and Lithuania, 8 percent in Latvia. In Estonia and Latvia, where negative population trend was partially offset by increasing share of working age population, the driving force was falling participation rate, but in Lithuania declining population was the major factor. During this period employment in Estonia and Lithuania has shrunk by 7 to 8 percent, of which over a half was due to rising unemployment rates, while contribution from the contraction of the labour force was about 3 percentage points. In Latvia falling unemployment has almost completely offset the effect of labour force contraction.

During the final episode (between 2000 or 2001 and 2003) employment growth was explained by falling unemployment rates completely in Estonia and Lithuania and by a major part in Latvia. Only Latvian labour force has changed significantly (by 2.6 percent, despite falling population), thanks to increase in participation and share of working age population.

Evolution of employed, unemployed and inactive population in each country is displayed in Figures 4 and 5. This time all indicators are in thousand, allowing for an accurate year-by-year balance. A detailed analysis will follow shortly, but one observation is hard to miss: For the core working age group, 15-64, the healthiest trends – declining unemployment and inactivity accompanied by growing employment, indicating rather flexible labour market, are found in Latvia in 2001-2003.

The early transition data are available only for Estonia. In each of the years 1990-1993 a substantial part of displaced workers in Estonia went to inactivity (Figure 4, middle panel). This might suggest an incidence of discouraged worker effect. Inspection of inactivity reasons reported by LFS respondents confirms that number of discouraged workers[3] increased by some 10 thousand between 1989 and 1993, but total increase of inactivity was 65 thousand. Early retirement, ageing, and disability were major contributors (Table 5). Vork and Habicht (2001) suggest that rules for granting disability were eventually relaxed to enable displaced workers to cope. In 1994 – 1997 fall in employment was almost completely (except for some 5 thousand persons in 1995) balanced by growth of unemployment and emigration, and increase in stock of discouraged workers slowed down. Number of disabled continued to increase.

Inspection of the labour market dynamics in 1998-2003 reveals that decrease in employment during the recession caused by Russian financial crisis (1998-1999 in Estonia, 1998-2000 in Latvia, 2000-2001 in Lithuania), as well as later decrease in unemployment in 2001-2002 in Estonia was partially absorbed by inactivity (Figures 4 and 5, middle panels). Number of discouraged workers went up. But discouragement was not the major factor. Inactivity growth was driven by sharply increasing number of students among the youth, which was partially offset[4] by decreasing number of pensioners (see Table 5; Table 6 provides the schedules of changes in statutory retirement age in the three countries). Increasing trend in the stock of discouraged workers was stopped in the last years of observation (2001-2003 in Latvia, 2002-2003 in Estonia and Lithuania), when employment went up in all three countries. The patterns of change were different, however. In Latvia, both unemployment and inactivity (including discouragement) were significantly reduced. In Estonia, number of discouraged workers and unemployed dropped in 2002, when total inactivity increased because of students; in 2003 total inactivity declined, while discouragement and unemployment did not change much. In Lithuania, unemployment and discouragement went down but total inactivity was not affected.

Proportion of inactive persons, who have not started job search because they do not know how and where to search, has been steadily decreasing in Latvia, indicating gradual

[3] Here the term "discouraged worker" is used loosely, referring only to the reported reason for not seeking a job. According to the standard definition, only those inactive persons, who would like to work and are available for work, are categorized as discouraged. See further sections for a more detailed discussion of discouraged worker effect in the Baltic countries.
[4] Except for the years 1999-2001 in Latvia.

improvement in the functioning of the labour market (this indicator is not available for the other two countries).

Labour market flows

A better understanding of labour market dynamics can be gained by analysing probabilities of transition between employment, unemployment and inactivity. Figure 6 displays recent history of transition probabilities for each of the three Baltic countries: 1997-2001 for Estonia, 1997-2002 for Latvia, and 1999-2003 for Lithuania. EU-15 data for 1997-98 and 1995-96 (European Commission, 2002, Table 22) will be used for comparison. The discussion here will focus on flows from and to inactivity.

About 4 percent of employed leave labour force every year in Estonia and Latvia; EU-15 figure was somewhat higher, close to 5 percent. Temporary increase of outflow from employment to inactivity observed in Latvia between 1999 and 2000 can be attributed to the already mentioned cap on pension benefits for working pensioners. In Lithuania annual outflow was significantly higher, about 6 percent, in 1999-2001, but dropped to 3 percent in the last two years, following acceleration of the pension reform (see Table 6).

Outflow from unemployment to inactivity can be thought of as related to discouraged worker effect. In Latvia annual rate of this outflow in 1997-2001 was fluctuating around 20 percent, comparable to EU-15 level of 17-19 percent; however, in 2002 the estimated outflow increased to 25 percent. In Lithuania rate of transition from unemployment to inactivity has decreased from 18 percent in 1999-2000 to 12-13 percent in the last two years of observation. In Estonia incidence of discouragement, according to this measure, was very low in 1997-2000[5] but jumped to a level similar to Lithuania (14 percent) between 2000 and 2001.Qualifying these changes one has to take into account that for the last year of observation in Estonia and Latvia, and for the last two years in Lithuania, transition rates are based on the retrospective question, which have a tendency to classify some of the last year's inactive as unemployed, thus overestimating the outflow from unemployment to inactivity (previous estimates are based on matching sub-samples). Decrease of the outflow in Lithuania, however, cannot be attributed to change in methodology (moreover, for 2002-2003 this outflow is even smaller, 10.6 percent, when estimated over the matching sub-sample). Interestingly, rate of transition from unemployment to employment in the Baltic countries has

[5] One cannot exclude that status in January as the base for calculations, in contrast with 2nd quarter in other countries, resulted in an underestimation of Estonian outflow in 1997-2000.

been very much the same as in EU-15 (around 30 percent), despite much higher unemployment rate.

Transitions from inactivity to either unemployment or employment are indicative of increasing labour force participation. Recent rates of outflow to unemployment (3 to 4 percent in most cases) and to employment (around 6 percent in Estonia and Latvia) are somewhat above the ones found in EU-15. In Lithuania the latter rate was slightly lower (4 percent). In all three countries overall rate of transition from inactivity to labour force features increasing trend in the last two years of observation.

Age and gender dimensions of participation

Next we turn to age and gender dimensions of labour force participation. Table 8 provides the data.

Baltic teenagers of both genders are much less likely to participate in the labour force than their counterparts in EU 15. Participation rates of 15-19 years old, which in 1997-98 were around 25 percent for males and around 20 (14 for Lithuania) percent for females, by 2003 dropped to 15-16 percent for males in Estonia and Latvia, 9-11 percent for males in Lithuania and females in Estonia and Latvia, and just 6 percent for Lithuanian female teens. In EU 15 these rates were stable at 31 to 33 percent for male teenagers and at 25 to 27 percent for their female counterparts. Plausibly, recent fall in Baltic teenagers' participation is related to growing real income of their parents. Late entry into the labour market is of course a consequence of high participation in education, but as OECD (2003a) suggests, it may also indicate a shortage of temporary and part-time jobs of the type that would be suitable for combining with studies in secondary school. Unlike the United States, the United Kingdom and much of northern Europe, there is also no strong tradition for teenagers to work.

Activity rates of 20 to 24 year olds in the Baltic countries have also decreased since 1997-98, especially strongly in Lithuania. Females of this age in all three countries, as well as young males in Lithuania, have participation rates well below the average level of EU 15, which was not the case in 1998. As discussed above, education is the main reason of inactivity of this age group. However, Latvia, where tertiary enrollment rate was as high as in Lithuania and above the Estonian level, featured substantially higher youth labour force participation rates. Gender gap in participation of the youth in the Baltic countries is larger than in the EU 15, because females here are more likely to continue education than males.

In the prime age group, 25 to 54, all three countries by 1998 had men's activity rate very close to the EU 15 average, while women's participation was by 12 to 17 points higher in the Baltic countries. Five years later, Baltic prime age men's activity has slightly decreased and was 1.5 to 3.0 points below the EU 15 level, while women's participation was 8 to 13 points above the EU 15 average (the latter has gone up by 4 points).

Activity rate of men aged 55 to 59 has decreased somewhat since 1997-98 in Estonia and Latvia; yet it is slightly above EU 15 average in Latvia and substantially above this level in Estonia and especially Lithuania. Due to pension reforms in the Baltic countries (see Table 6), participation of women aged 55-59, as well as of men aged 60-64, has been growing much faster than in EU 15. By 2003, activity rate of Baltic women aged 55-59 was 10 to 16 percentage points above the EU 15 average. This is a remarkable development, given that in 1998 Latvia was 4 points behind EU 15, and Lithuania was just one point above. Likewise, in Latvia and Lithuania, activity rate of men aged 60 to 64 in 2003 was 5 to 7 points above EU 15 average, while in Estonia, where the pension reform has started earlier and provides the largest incentives to deter retirement, this rate was 17 points above EU 15 level[6].

Baltic females aged 60 to 64 are still eligible for retirement, yet their participation rates are on the rise and in 2003 were substantially above the EU 15 average, especially so in Estonia (almost 20 points difference).

Overall, activity rates of the 15-64 age group in the Baltic countries are some four to five points below the average EU 15 level for men and three to five points above it for females. Resulting activity rate for both genders in 2003 was just below the 70 percent level of EU 15. Gender gap in participation in the prime age, as well as for 55-64 years old (except 60-64 in Lithuania), is smaller in the Baltic countries than it is in EU 15.

As far as elderly are concerned, Estonian case suggests strongly that this age group can become a real asset in the labour market: after introduction, in 1996, of the possibility to receive old-age pension simultaneously with labour income, the number of economically active individuals aged 65 to 74, which was falling in the early years of transition, started to rise and almost doubled by 2003, while number of inactive persons has stayed constant (Figure 4, lower panel).

The fact that income elasticity of supply is high for those in pre-retirement and retirement age[7] is confirmed also by Latvian and Lithuanian experience. In Latvia, elderly labour force has contracted by 19 percent in 2000, when restrictions on pensions for working retirees were introduced, but when the restrictions were eliminated by the Constitutional Court in 2002, number of economically active persons aged 65 to 74 returned to the previous level (Figure 5, lower panel); activity rates of men aged 60 to 64 and women aged 55 to 59

[6] Pensions are enhanced by 10.8% per year of postponed retirement in Estonia and by 8% in Lithuania; in Latvia the NDC system also ensures that workers benefit from postponed retirement.

[7] See Prescott (2004) for recent evidence on high elasticity of labour supply in G7 countries.

have also increased sharply, by 7 and 11 percentage points respectively, in 2002 (Table 8). In Lithuania, targeted (and somewhat higher than ordinary) unemployment benefits were introduced in 2002 for persons who will reach statutory retirement age in 5 years or less (this is the main reason behind increase in average UB in 2002-2003 reflected in Figure 2). On top of this, after-tax minimum wage went up by 4.4 percent in 2002 and by 5 percent in 2003 (Table 3). These developments clearly contributed to rise in activity rate of women aged 55 to 59 by 11.6 percentage points in 2002-2003 (while retirement age increased just by 6 months per annum, same as in 2001 and only by 2 months more than in 1998-2000).

In Lithuania, pensions are reduced when recipients have work income. Persons earning more than 1.5 times the minimum wage receive only basic pensions. With lower earnings, the supplementary pension is reduced if the earnings exceed the minimum wage (OECD 2003a). On the other hand, average pension benefits in Lithuania are somewhat higher than in Estonia and Latvia relative to average wage, while after-tax minimum wage exceeds average pension only in Lithuania (Table 6). This suggests that those Lithuanian elderly, who are not prepared to accept unskilled jobs with minimum wage, have less work incentives than their Estonian and Latvian counterparts. Indeed, labour force participation of the 65-74 years old in Estonia has reached 16 percent in 2003, while it was 12 percent in Latvia and less than 8 percent in Lithuania; moreover, in Lithuania less than a half of employed in this group were wage earners, while in Estonia this proportion was above three quarters.

Determinants of participation

Table 9 presents results of panel estimates (population averaged probit, assuming equal error correlation within panels) of labour force participation of population aged 15-74 by gender, based on recent labour force surveys in Estonia and Lithuania. For Estonia we have used 2001 LFS. Initially there were one or two observations for each respondent, but due to very detailed retrospective part it was possible to track all necessary variables back to January 2000 with quarterly intervals, so we end up with more than 55 thousand observations, average panel size is about 6. For Lithuania we have used 2nd and 4th quarters of two consecutive years, 2002 and 2003, with about 39 thousand observations; some respondents are observed twice and some once, so average panel size is 1.6.

Basic controls include education (6 categories), 5-year age groups, ethnicity, marital status, dummies for having one or more children, residence in rural area, and region[8] fixed effects. To capture effect of minimum wages, as well as of average wage growth and local economic conditions, we include one or two of the following macro-level trends: real minimum wage at the beginning of the quarter, last quarter's real national average wage and last quarter's unemployment rate, as well as region-specific last year's real average wage and last year's local unemployment rate (all these variables in logarithmic form; for Estonia quarterly county level wage data were used). Interactions of young and/or old age dummies with wage variables are included when relevant.

To account for the coordination of the labour supply decisions within the household we include spouse's or partner's wage (set to zero for singles), and interactions of young age dummies with parents' wage (set to zero for persons not living with parents). These measures of non-labour income are divided by the number of relevant core family members: spouse's wage by 2 plus number of children under 15; parent's wage by number of parents plus number of their children (in this household) under 20 or 25.

It turns out, however, that for women in both countries, as well as for men in Lithuania, partner's wage is extremely insignificant determinant of participation (Table 10). This is typical situation for transition countries (see e.g. Saget, 1999). Estonian men are significantly more likely to participate if their wives earn more, likely through correlation of partners' educational attainment. Parents' earnings effect has expected negative sign for people younger than 25, but is significant only for Lithuanian young females. Therefore in the baseline model we do not use non-labour income. In this model we also do not control for being a pupil or student (effects of including this variable are discussed later).

Comparison of the results reveals that other things equal, young and old age participation gaps for both genders (except female teenagers) are substantially wider in Lithuania than in Estonia. On top of this, young Estonians, as well as Lithuanian female teens have 5 to 10 points higher participation rates when there is no prime age persons in the household; surprisingly, for Lithuanian females aged 20-24 this effect has opposite sign, perhaps indicating that many of them live separately and receive external financial support.

Higher education, as well as vocational (without secondary) education has a much stronger effect on men's participation in Estonia than in Lithuania, but for women it goes the

[8] In Estonia we use 15 counties, but the capital city (400 thousand population) is separated from the rest of respective county; excluding capital city, average population these units is about 90 thousand. In Lithuania we use fixed effects for 10 counties and three large cities, but local wages and (registered) unemployment are measured at municipality level; there are 60 municipalities with average population 59 thousand.

other way around. For both genders, postsecondary professional education boosts participation stronger in Estonia.

In both countries women belonging to ethnic minorities, have 5 to 6 percentage points lower participation rates than their otherwise similar majority counterparts (situation is not different in Latvia, see Hazans 2004b). For men the ethnic participation gap is not significant; however, interestingly enough, it becomes significant when controls for being a student are included (see Table 10). An explanation comes from the following equation, where LF is labour force,

$$\Pr(LF)=\Pr(Student)\,\Pr(LF|Student) + \Pr(Non\text{-}Student)\,\Pr(LF|Non\text{-}Student) \qquad (1)$$

Hence, denoting ethnic Lithuanians with subscript 1, minorities with 2, the difference between the two with Δ, and abbreviating Student as S, one has (conditional on characteristics):

$$\Delta \Pr_1(LF) = \Delta \Pr(S)\Pr_1(LF|S)+ [\Delta \Pr(LF|S)]\, \Pr_2(LF|S)+ \Delta \Pr(NS)\, \Pr_1(LF|NS)$$

$$+[\Delta \Pr(LF|NS)]\Pr_2(LF|NS)$$

$$= \Delta \Pr(S)[\, \Pr_1(LF|S)- \Pr_1(LF|NS)] + [\Delta \Pr(LF|S)]\, \Pr_2(LF|S) +[\Delta \Pr(LF|NS)]\Pr_2(LF|NS).$$

Probability to be a student is smaller for minorities in Lithuania (in our sample 0.092 and 0.109 respectively), and of course $\Pr_1(LF|S)=0.193 < \Pr_1(LF|NS)=0.731$. So the first term on the RHS is negative, while the second and the third are positive according to Table 10 and the total result in not significantly different from zero. Other things equal, minority males are less likely to be in the labour force conditionally on studying or not studying, but this is compensated by being more likely in a group with higher participation.

Having children decreases activity of Estonian females a lot more strongly than their Lithuanian counterparts.

Ceteris paribus rural – urban participation gap is minus 4 percentage points for Estonian men, while it is plus 6 points for Lithuanian men.

Finally we turn to minimum wage and local economic conditions. According to the standard economic theory (Ehrenberg and Smith, 2003) rising minimum wage increases participation. But on the other hand, it negatively affects demand for labour, and hence, through discouraged worker effect can adversely influence participation. In Lithuania, increasing after-tax real minimum wage appears to have, on average, positive effect on participation. Reported marginal effect implies that a modest 5 percent increase in after-tax minimum wage results in 2.7 percentage points higher participation for women and 1.2 points for men. In Estonia, by contrast, positive effect of minimum wage on participation is found only for teenagers of both genders and for young males. A 10 percent increase in real minimum wage boosts participation of these two groups by two and three percentage points respectively. For other groups estimated effect is negative. This is likely to be related to negative effect of increased minimum wages on labour demand for low skilled, which was found in Hinnosaar and Room, 2003 (our model controls for labour demand only indirectly, through unemployment).

Wage growth differentials between regions appear to have, on average, no significant effect on participation in Lithuania. In Estonia, female teenagers and older females are more likely to participate when average wages are higher or in the regions with higher wage growth, while for women aged 20-24 there is an opposite effect (in contrast with Lithuania, the respective variable is measured quarterly and varies over time independently from minimum wage; when national trend and deviation are included, both have positive signs; reported results refer to a model where these two effects are not disentangled).

In regions with higher unemployment, males in both countries, as well as females in Estonia are less likely to participate in the labour market: if unemployment rate doubles, other things equal, activity rate goes down by about 2 percentage points (3 points for Estonian females). This is indicative of *discouraged worker* effect. For Lithuanian females, by contrast, the effect has opposite sign (and same magnitude), suggesting that *added worker* effect is at work.

After accounting for minimum wage, there is no significant time trend in participation (although there is a very strong seasonal effect in Lithuania: participation is 4 to 5 points higher in the second quarter than it is in the fourth, likely due to tourism).

Table 10 reports the results with controls for non-labour income and studies (the original LFS samples, without the retrospective extensions, are used for both countries). In both countries non-student males aged 20-24, who are not living together with wage-earning parents, are as likely to be labour force members as otherwise similar males aged 40-44. However, when being a student is controlled, parents' wages tend to increase labour force participation of young males in Estonia, while in Lithuania an opposite effect is observed[9]. While each of the effects is not significant even at 10 percent level, the difference between the countries is. Parental wage effect is virtually absent for females aged 20-24 in both countries, but this because it works through participation in education.

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[9] Dummy for the 20-24 age group is interacted with deviation of parental income per core family member from its mean value, standartised by national average net wage. Using deviation ensures that interaction does not distort the main effect of the age group dummy. Recall that

In Lithuania non-students females aged 20-24 are relatively a lot more active: just 9 points behind the 40-44 years old, as opposed to 30 points in Estonia. For female students of this age, however, the participation gap is 61 percentage points, while it is just 45 points in Estonia. For male students aged 20-24 in both countries probability to participate in the labour force is 61 to 63 points lower than for otherwise similar males aged 40-44. But on top of this there is a negative effect of being single: minus 15 points for Lithuanian males, and minus 5 points for their Estonian counterparts.

Partner’s wage has negative (though not significant) effect on participation only for Estonian women.

As mentioned before, controlling for studies makes the ethnic participation gap larger. Females of non-Estonian ethnicity are 10 percentage points less likely to be in the labour force than otherwise similar ethnic Estonian females; in Lithuania this gap is 7 points, but for males it is two times smaller than for females and significant only at 10 percent level (for females the effect is very significant).

Discouraged workers: a closer look

In section three above we have discussed the dynamics of inactive persons who reported discouragement as the reason why they do not look for a job. Strictly speaking, according to the standard definition, only those who nevertheless would like to work and are available for work, are categorised as discouraged workers. A relaxed definition includes all inactive persons who would like to work and are available for work, disregarding the reason for not seeking a job. Discouraged workers can be viewed as the immediate reserve of the labour force.

Proportion of discouraged workers (relaxed definition) among inactive population aged 15 to 74 is quite high in Estonia: 17 percent for males and 13 percent for females (year 2001 data) and even higher in Latvia (24 percent for males and 20 percent for females in 2002). In Lithuania (2002-2003), 5 percent of inactive men and 4 percent of inactive women aged 15-74 fall into this category. With respect to total population aged 15 to 74 the difference between the two countries is smaller: 5.4 percent of this age group in Estonia (2001) were discouraged workers in the broad sense, a 7 times higher proportion than in the beginning of 1999 (this trend is consistent with Table 5 data on reasons for nor seeking a job). In Lithuania this proportion has increased from 3.1 to 5.7 percent between 2000 and 2002, but felt to 3.8 percent in 2003.

Table 11 reports probit estimates of determinants of discouragement among inactive population aged 15 to 74 in Estonia and Lithuania. Conditional on inactivity, the probability of being discouraged (that is, being ready to take on a job) peaks at 41-42 years of age for females in both countries and males in Estonia, while for Lithuanian males it is maximal at 31 years of age.

Other things equal, females with secondary general and secondary vocational education in both countries, as well as with vocational (without secondary) education in Estonia are most likely to be discouraged. For inactive males in Estonia education does not affect likelihood of being discouraged, while in Lithuania inactive males with vocational (without basic) education are most likely to be available for work, followed by the ones with professional or vocational secondary education. Students and schoolchildren are significantly less likely to be ready for a job than otherwise similar inactive persons who are not studying, but in Lithuania this effect is less pronounced than in Estonia.

In Lithuania, inactive females with one child are more likely to be available for work than childless women, other things equal.

Ethnicity of inactive person in Lithuania does not have a significant effect on likelihood to be a discouraged worker (despite the fact that the proportion of discouraged workers among minorities was 5.7 percent and just 4 percent among ethnic Lithuanians; these are proportions out of inactive population aged 15 – 74, average for 2002-2003). In Estonia LFS provides information on state language skills, which reveals that inactive males who do not speak Estonian language, and especially those who do not even understand it, are most likely to be discouraged. For females this effect is not found. By contrast, in Latvia, inactive females belonging to ethnic minorities are more likely to be discouraged than otherwise similar ethnic Latvian females (Latvian results are available on request).

There is evidence that increasing real after-tax minimum wage in Lithuania has had a reducing effect on discouragement in Lithuania, especially for women (a 10 percent increase in minimum wage reduces likelihood of discouragement by one percentage point). Inactive persons in Lithuania, especially if they are young, are less likely to be discouraged when they live in a municipality with higher average wages, but the size of this effect is small.

Local unemployment does not manifest itself as a factor influencing discouragement in Tale 10, but this is because region fixed effects are included, while small panel size (one to two observations) does not allow for the variation over time to play a role. In alternative models without region fixed effects, or with narrow definition of discouragement (in which case we have longer panels), local unemployment in Estonia has a strong positive effect on discouragement for both genders. In Lithuania it is not the case; the effect is positive as well but not significant even when both county dummies and local wages are removed from the model. These results are available on request.

Conclusions

From the labour market perspective, the transition period in the Baltic countries can be broken down into three episodes. Similarly to Hungary, Bulgaria, and Czech R., labour force contraction was responsible for at least half (in Estonia even two thirds) of the massive employment reduction between 1989 and 1996/7; however, in the Baltic case demographic trends were much more important as the reason behind declining labour force, especially in Latvia and Estonia. During next three or four years falling participation rates in Latvia and Estonia and declining population in Lithuania caused a further labour force reduction, although at a slower path. During this period employment in Estonia and Lithuania has shrunk by 7 to 8 percent, of which over a half was due to rising unemployment rates, while in Latvia falling unemployment has almost completely offset the effect of labour force contraction. During the final episode (between 2000 or 2001 and 2003) employment growth was explained by falling unemployment rates completely in Estonia and Lithuania and by a major part in Latvia. Only Latvian labour force has somewhat increased, despite falling population.

The discouraged worker effect has been at work in all three countries, although the dynamics of discouragement was not always consistent with trends in participation, which were largely defined by the pension reforms, changes in regulations related to working pensioners, and increasing enrolment of the youth in further education. In Lithuania, there is also a recent evidence for added worker effect in districts with higher unemployment.

In all three Baltic countries, recent rates of transition from unemployment to employment and to inactivity were similar to those found in EU-15, while overall rate of transition from inactivity to labour force features increasing trend in the last two years of observation.

A dramatic decrease in youth participation rates and sharp increase in participation of females aged 55 to 59, as well as 60 to 64 years old men and women, took place in all three countries between 1997-8 and 2003. However, large differences between recent rates across countries suggest that there is substantial room to increase labour supply. The following recommendations are based on comparison of age- and gender-specific activity rates, as well as on the econometric analysis of labour force participation, which controls for also for factors other than age and gender.

First, higher labour force participation by the teenagers, as well as students aged 20 to 24 in Lithuania and in Estonia, could be pursued; in Estonia, this applies also to non-student females aged 20 to 24. Second, higher participation is a realistic option for Latvian females approaching retirement age, whose activity rate is currently 5 points below the level found in the other two countries. Another possibility is mobilisation of both men and women in their early 60s in Latvia and Lithuania, where participation rates in this age group are at least 10 points below those found in Estonia (although higher than in EU 15). Finally, Lithuanian population aged 65 and older has substantially lower participation (especially in paid employment) than Estonian population of the same age; moreover, activity of this group is stagnant in Lithuania, while it is rising in Estonia.

Why are the older segments of population in Estonia more active than in the other two Baltic countries? As far as Lithuania is concerned, the restrictions on pensions for working pensioners clearly play a role. While such restrictions are now removed in Latvia, they are likely to have a lasting effect as well, because it is more difficult for an older person to re-enter labour market. On the other hand, postponed retirement in Estonia enhances pensions stronger than it does in Lithuania. Perhaps one more reason is that average wages in Estonia are higher than in Latvia and Lithuania both absolutely and when compared to average pension (Table 7); this also true for minimum wages when Estonia and Lithuania are compared.

In all three Baltic countries, representatives of ethnic minorities (especially females) are significantly less likely to be labour force members than their majority counterparts; closing this gap will substantially increase overall participation rates.

Based on participation effects, it appears that postsecondary professional education better suits labour market needs in Estonia than it does in Lithuania; the same is true for higher and vocational (without secondary) education for men, while for women two latter types of education boosts participation stronger in Lithuania.

Increasing after-tax real minimum wage appears to have, on average, positive effect on participation in Lithuania, while in Estonia such effect is found only for teenagers of both genders and for young males. Targeted unemployment benefits seem to raise participation of pre-retirement age persons in Lithuania.

Significant portions of inactive population aged 15 to 74 in all three Baltic countries are not engaged in job search, although they are willing to work and available for work. When considered against total (rather than inactive) population of this age, this group, which can be seen as the immediate reserve of the labour force, represents more than 5 percent in Estonia, about 4 percent in Lithuania, and 8 percent in Latvia. Given that inactive persons are most likely to fall into this category (broadly defined discouraged workers) when they are around 40 years of age (even 30 for Lithuanian men), this is a real reserve. In Estonia, inactive males who do not speak Estonian language, and especially those who do not even understand it, are most likely to be discouraged. In Latvia, inactive females belonging to ethnic minorities are more likely to be discouraged than otherwise similar ethnic Latvian females.

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Table 1 Educational attainment of adult population and enrolment into further education of the youth in the EU-15 and selected countries of Central and Eastern Europe, 2002

	EU-15	ACC-12	SI	BG	HU	RO	EE	LV	LT
Education	**Percent distribution of population aged 25-64 by highest level of completed education**								
Basic or less	35.4	19.3	23.2	28.5	28.6	28.9	12.5	17.4	15.2
Upper secondary	42.9	66.2	62.1	50.4	57.3	61.1	57.9	63.1	63.3
Tertiary	21.8	14.5	14.8	21.1	14.1	10.0	29.6	19.6	22.5
	Enrolment in further education of population aged 18-24 with basic education or less								
	81.2	91.3	95.2	79.0	87.7	76.8	87.4	80.5	85.7

Notes: ACC-12 – average for the 10 new EU members, Bulgaria and Romania. Country abbreviations: SI - Slovenia, BG – Bulgaria, HU – Hungary, RO – Romania, EE – Estonia, LV – Latvia, LT – Lithuania.
Source: Franco and Blondal (2003).

Table 2 Immigrants from new EU member states registered in UK, May – October 2004 (per 1000 population of the sending country)

LT	LV	SK	PL	EE	CZ	HU
4.6	2.8	1.7	1.6	1.3	0.6	0.3

Notes: See Notes to Table 1 for country abbreviations.
Source: UK Home Office and own calculation.

Table 3 Minimum wage developments in the Baltic countries

	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001	2002	2003
Minimum wage – average wage ratio (percent, annual average)												
Estonia	36.4	21.1	20.2	18.9	22.8	23.6	26.7	28.2	28.5	29.0	30.1	32.1
Latvia	27.4	26.6	30.6	31.3	36.0	31.7	31.5	35.5	33.4	34.6	34.7	36.4
Lithuania	24.2	19.7	17.4	28.0	38.8	48.1	44.9	43.6	44.3	43.8	42.4	40.7
Nominal increase during the year (December on December)												
Estonia		50.0	50.0	0.0	51.1	24.3	30.2	13.6	12.0	14.3	15.6	16.8
Latvia	226	100	86.7	0.0	35.7	0.0	10.5	19.0	0.0	20.0	0.0	16.7
Lithuania	240	182	35.4	177	66.7	33.3	7.5	0.0	0.0	0.0	0.0	0.0
Number of changes during the year												
Estonia		1	1	0	1	1	1	1	1	1	1	1
Latvia	4	1	2	0	1	0	1	1	0	1	0	1
Lithuania	5	10	4	5	2	2	1	0	0	0	0	0
Real increase in after tax[a] minimum wage during the year (December on December)												
Estonia	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	19.3	9.4	6.6	9.7	12.6	15.6
Latvia	-69.2	48.3	47.8	-18.8	20.0	-6.5	7.5	15.4	-1.8	16.3	-1.4	12.6
Lithuania	-73.1	-2.2	-6.7	104.1	38.7	32.0	5.9	-0.3	-1.4	-2.0	4.4	5.0

Note: [a] For Estonia – gross minimum wage. *Sources:* National Statistical offices and own calculation.

Table 4 Break-down [a] of the changes in economically active and employed population

Percent

	Estonia			Latvia			Lithuania		
First year	1989	1997	2000	1989	1996	2000	1989	1997	2001
Last year	1997	2000	2003	1996	2000	2003	1997	2001	2003
Change in Employment	-26.8	-7.2	3.8	-32.7	-0.7	7.2	-22.8	-8.3	6.4
of which due to:									
Change in unemployment rate[b]	-9.2	-4.3	4.1	-17.9	7.9	4.5	-12.4	-5.6	6.0
Change in Labour Force	-19.4	-3.0	-0.3	-18.0	-8.0	2.6	-11.5	-2.8	0.4
of which due to:									
Change in Population	-10.2	-2.4	-1.2	-7.4	-3.6	-2.1	-2.4	-2.8	-0.7
Change in working age population %	-0.7	1.4	1.2	-1.5	2.2	1.5	-1.4	0.6	1.2
Change in participation[c]	-9.7	-2.0	-0.3	-10.3	-6.6	3.3	-8.2	-0.6	-0.1
	Activity, unemployment and employment rates, age 15-64								
Activity rate, first year	78.9	72.3	70.4	81.9	71.7	67.2	77.6	70.1	69.4
Activity rate, last year	72.3	70.4	69.8	71.7	67.2	68.6	70.1	69.4	69.7
Unemployment rate, first year	0.5	9.3	12.8	0.0	20.5	14.6	0.0	12.6	17.6
Unemployment rate, last year	9.3	12.8	10.2	20.5	14.6	10.0	12.6	17.6	12.5
Employment rate, first year	78.5	65.2	60.7	81.9	57.0	57.3	77.6	61.3	57.2
Employment rate, last year	65.2	60.7	62.6	57.0	57.3	61.8	61.3	57.2	60.9

Notes: [a] Numbers in the table are changes in percent rather than log points, hence totals are not exactly equal to the component sums. Demographic indicators refer to beginning of the years. Labour market indicators are annual average.

[b] Numbers in this row are percentage changes in 1- *u*, so they are negatively related to changes in unemployment rate. [c] Participation here is ratio of total labour force to working age population, so it differs slightly from labour force participation rate for the 15-64 age group.

Table 5 Inactive population by reason for not seeking a job

	Distribution, percent					Change vs. previous year, thsd					
Estonia, age 15-69	1989	1993	1997	2001	2003	1998	1999	2000	2001	2002	2003
Studies	34.5	27.0	26.2	31.6	36.1	6.7	9.6	-0.5	4.4	21.2	-6.1
Retirement	35.9	41.9	40.6	32.9	29.4	-1.9	-4.2	-8.6	-7.4	-15.1	4.0
Disability	8.9	9.7	12.7	13.1	13.3	-1.0	0.8	4.0	-1.5	3.0	-2.2
Discouragement	**0.6**	**3.3**	**4.8**	**6.8**	**5.5**	**1.9**	**1.7**	**0.2**	**3.3**	**-4.7**	**0.4**
Family & personal	16.0	14.2	12.3	11.2	11.3	-0.8	-0.7	-1.7	0.7	6.7	-6.3
Other	4.0	3.9	3.5	4.4	4.4	-1.6	-0.3	2.5	2.7	-3.1	3.1
Total	100.0	100.0	100.0	100.0	100.0	3.3	6.9	-4.1	2.2	8.0	-7.1
Latvia, age 15-64	**1996**	**1997**	**2000**	**2001**	**2003**	**1998**	**1999**	**2000**	**2001**	**2002**	**2003**
Studies	27.8	28.8	35.1	37.4	39.8	37.4	-18.9	27.7	7.4	-4.6	7.6
Retirement	29.0	32.0	30.8	32.2	21.4	-8.0	11.9	5.2	3.0	-48.8	-10.0
Disability	12.1	10.8	9.3	9.3	9.5	-2.7	7.4	-7.4	-1.1	-1.4	0.6
Discouragement	**7.3**	**9.6**	**9.6**	**8.3**	**8.1**	**-5.7**	**5.3**	**4.5**	**-7.5**	**5.0**	**-8.0**
Do not know where and how to seek	4.4	3.6	2.5	1.1	n.a.	-1.5	-1.3	-1.0	-7.2	n.a.	n.a.
Family & personal	10.1	9.2	8.7	6.5	14.2	-6.2	6.2	2.0	-12.0	30.3	5.0
Other	9.3	6.1	4.1	5.2	7.0	-6.8	7.4	-7.9	4.9	9.8	-2.1
Total	100.0	100.0	100.0	100.0	100.0	6.5	17.9	23.3	-12.5	-15.3	-7.0
Lithuania, age 15-64		2000	2001	2002	2003				2001	2002	2003
Studies		41.7	43.2	46.0	49.6				17.0	23.9	22.4
Retirement		24.8	23.0	20.5	18.5				-8.1	-15.6	-15.0
Disability		11.6	14.0	15.2	15.5				18.5	9.7	1.2
Discouragement		**6.8**	**5.9**	**5.2**	**3.9**				**-5.3**	**-4.2**	**-9.3**
Family & personal		8.3	8.2	8.3	7.8				0.6	1.6	-3.9
Other		6.8	5.8	4.8	4.7				-6.1	-6.1	-1.0
Total		100	100.0	100.0	100				16.5	9.2	-5.5

Sources: Estonia – Statistical Office of Estonia (www.stat.ee); Latvia and Lithuania – calculation based on LFS data.

Table 6 Statutory retirement age

	1989	1993	1994	1997	2000	2001	2002	2003
Estonia								
Men	60	60.5	60.5	61.5	62.5	63	63	63
Women	55	55.5	55.5	56.5	57.5	58	58.5	58.5
Latvia								
Men	60	60	60	60	60.5	61	61.5	62
Women	55	55	55	56.5	58	58.5	59	59.5
Lithuania								
Men	60	60	60	60.5	61	61.5	62	62.5
Women	55	55	55	56	57	57.5	58	58.5

Note: In Latvia changes for women in force since July 1 of corresponding year. Intermediate steps in 1995-96 and 1998-99 not shown. *Source*: National Ministries of Welfare.

Table 7 Average old-age pensions as per cent of average and minimum wages

	1993	1994	1995	1996	1997	1998	1999	2000	2001	2002	2003
Estonia											
Av. pension/Av. gross wage	22	22	26	31	31	29	35	32	28	26	27
Av. pension/Av. net wage	27	27	32	40	40	38	45	41	36	34	36
Av. pension/ Minimum wage after tax			na	na	na	na	na	118	110	99	99
Latvia											
Av. pension/Av. gross wage	47	43	40	41	39	42	43	40	39	37	35
Av. pension/Av. net wage	54	51	49	52	53	58	58	55	54	52	49
Av. pension/ Minimum wage after tax			163	143	163	178	161	161	136	143	131
Lithuania											
Av. pension/Av. gross wage			31	33	32	32	32	33	32	32	32
Av. pension/Av. net wage			41	43	43	43	44	46	45	44	44
Av. pension/ Minimum wage after tax			109	93	79	84	89	90	89	89	90

Source: National Statistical offices and own calculation.

Table 8 Labour force participation rates, 1997-2003 (annual average)

	Men							Women						
	1997	1998	1999	2000	2001	2002	2003	1997	1998	1999	2000	2001	2002	2003
Age 15 to 19														
Estonia	24.8	21.3	15.4	16.6	16.8	10.9	15.1	19.2	15.8	12.4	16.0	14.1	6.4	8.9
Latvia	27.0	23.5	21.9	16.4	14.1	18.8	15.9	20.1	15.8	12.8	9.8	9.6	11.5	10.9
Lithuania		23.5	21.3	16.2	11.1	8.6	8.8		13.5	14.3	6.8	6.1	4.9	6.0
EU 15	31.5	32.2	32.8	33.1	31.9	31.0	30.7	25.4	25.9	26.8	27.5	26.6	25.7	25.7
Age 20 to 24														
Estonia	79.5	79.1	78	80.4	78.3	70.5	71.2	58.3	62.3	57.6	56.5	54.8	51.3	53.9
Latvia	80.6	77.3	78.4	74.1	73.3	73.2	76.8	62.9	63.7	59.4	55	56.7	58.8	57.2
Lithuania		77.0	75.3	70.1	67.0	64.5	63.0		58.4	60.3	56.4	52.1	51.7	48.8
EU 15	69.7	69.6	69.7	69.8	69.1	68.8	69.8	58.8	59.3	59.6	59.9	59.0	59.1	59.8
Age 25 to 54														
Estonia	92.7	91.6	91.2	90.5	89.8	89.9	89.5	85.4	84.3	83.6	83.6	82.8	81.0	82.1
Latvia	89.7	91.4	90.3	88.0	89.4	89.2	89.6	83.7	83.2	82.5	83.4	83.5	82.4	82.9
Lithuania		92.1	90.6	89.7	90.1	90.8	90.4		87.4	89.2	87.9	88.0	87.4	87.1
EU 15	92.5	92.7	92.6	92.6	92.4	92.3	92.4	70.0	70.7	71.5	72.1	72.3	73.1	73.9
Age 55 to 59														
Estonia	78.5	76.9	74.1	76.0	74.6	72.4	75.2	52.3	54	52.8	52.3	56.9	67.7	65.3
Latvia	73.2	74.4	72.8	71.7	72.5	75.1	71.8	39.8	39.5	39.7	41.7	46.1	56.9	59.9
Lithuania		78.4	79.2	75.7	77.9	78.4	79.2		44.5	45.0	54.1	53.7	57.3	65.3
EU 15	69.7	70.0	70.0	70.0	70.3	71.4	71.4	42.6	43.4	44.4	45.5	46.4	48.0	49.4
Age 60 to 64														
Estonia	43.4	46.7	47.3	48.7	46.4	55.4	54.2	21.3	23.9	25.2	26.3	31.3	35.4	37.1
Latvia	36.8	32.5	34.2	35.9	34.1	41.3	42.1	20.8	17.6	18.5	17.6	22.0	23.8	26.9
Lithuania		35.8	36.6	38.3	39.6	40.3	44.5		16.2	17.6	18.0	14.2	17.5	20.5
EU 15	33.5	33.1	33.5	33.8	34.7	35.3	36.8	15.7	15.1	15.6	16.0	16.8	17.8	18.2
Age 15 to 64														
Estonia	78.5	77.4	76.0	76.1	75.2	74.1	74.5	66.7	66.5	65.2	65.3	65.4	64.3	65.5
Latvia	76.4	76.4	75.3	72.5	72.8	73.9	74.0	64.9	63.8	62.5	62.3	63.3	64.1	64.8
Lithuania		77.7	76.3	74.2	73.4	73.2	73.1		66.7	68.2	67.1	65.8	65.7	66.5
EU 15	78.2	78.4	78.5	78.6	78.4	78.4	78.6	58.1	58.7	59.5	60.1	60.3	60.9	61.6
	Men and women, age 15 to 64							**Men and women, age 65 to 74**						
Estonia	72.3	71.7	70.3	70.4	70.1	69.0	69.8	10.2	10.5	12.2	13.2	14.3	15.7	16.4
Latvia	70.4	69.8	68.6	67.2	67.9	68.8	69.8	12.0	11.3	12.4	10.2	9.9	12.6	11.6
Lithuania		72.0	72.1	70.5	69.4	69.3	69.7		9.3	8.4	10.3	8.3	6.9	7.8
EU 15	68.2	68.6	69.0	69.4	69.4	69.7	70.1	4.7	5.1	4.8	4.7	4.8	5.3	5.4

Sources: National statistical offices of Estonia, Latvia, and Lithuania; OECD.

Table 9 Determinants of labour force participation

	Estonia, 2000-2001				Lithuania, 2002-2003			
	Men		Women		Men		Women	
Mean Y	0.748		0.578		0.671		0.579	
	dy/dx	z	dy/dx	z	dy/dx	z	dy/dx	z
Education: Higher	0.289	10.62	0.322	9.4	0.215	13.00	0.370	19.77
Postsecondary professional	0.251	6.94	0.281	8.13	0.179	10.39	0.248	12.81
Secondary general	0.147	6.21	0.155	6.52	0.050	3.04	0.123	7.05
Secondary vocational	0.187	7.63	0.239	8.39	0.198	11.41	0.237	11.94
Vocational	0.190	7.30	0.095	2.56	0.108	2.09	0.192	2.11
Age 15-19	-0.358	-14.67	-0.477	-15.88	-0.647	-22.27	-0.724	-29.41
(Age 15-19) ×No prime age persons in the household	0.111	2.71	0.075	3.05	0.031	0.94	0.115	2.72
Age 20-24	-0.101	-6.87	-0.301	-11.13	-0.173	-8.22	-0.336	-16.45
(Age 20-24) × No prime age person in the household	0.070	5.57	0.048	1.93	-0.001	-0.06	-0.095	-2.24
Age 25 – 29	-0.013	-0.73	-0.165	-6.21	0.031	2.19	-0.069	-4.19
Age 30 – 34	0.018	1.06	-0.093	-3.56	0.029	2.04	-0.023	-1.65
Age 35 – 39	0.005	0.44	-0.008	-0.33	0.017	1.31	-0.027	-2.07
Age 45 – 49	0.006	0.56	-0.012	-1.2	-0.031	-2.17	-0.016	-1.18
Age 50 – 54	-0.026	-2.29	-0.091	-5.51	-0.062	-4.06	-0.051	-3.77
Age 55 – 59	-0.162	-8.14	-0.417	-14.99	-0.123	-7.09	-0.270	-13.92
Age 60 – 64	-0.436	-16.06	-0.580	-20.68	-0.452	-20.21	-0.648	-28.24
Age 65 – 74	-0.650	-18.8	-0.703	-25.6	-0.775	-31.9	-0.793	-36.16
Single	-0.074	-4.14	-0.100	-5.13	-0.165	-8.97	-0.062	-4.68
One child	0.105	8.16	-0.075	-2.33	0.071	2.99	-0.044	-2.36
More children	0.142	8.45	-0.202	-5.44	0.109	2.23	-0.103	-2.95
Ethnic minority	-0.009	-0.41	-0.064	-2.17	-0.022	-1.26	-0.048	-2.77
Disabled	-0.416	-6.96	-0.318	-5.84	n. a.	n. a.	n. a.	n. a.
Rural	-0.041	-2.51	-0.024	-1.27	0.059	4.16	0.005	0.35
MW=Log (min wage) (last Q): main effect	-0.141	-3.08	-0.079	-1.68	0.245	2.17	0.542	4.57
(MW-mean(MW))×(age 15- 19)	0.357	2.90	0.467	3.20				
(MW-mean(MW))×(age 20-24)	0.359	2.29						
AW=Log (avg. local wage) last year: main effect					-0.056	-0.94	-0.054	-0.81
(AW-mean(AW))×age 15-19			0.126	1.59				
(AW-mean(AW))×age 20-24			-0.146	-2.00				
(AW-mean(AW))×age 60-74			0.084	1.82				
Log (last year county unemployment rate)	-0.034	-1.32	-0.046	-1.74	-0.035	-1.56	0.033	1.44
County fixed effects (vs. capital city)	yes							
Min	-0.106	-2.60	-0.101	-2.01	-0.135	-3.26	-0.156	-3.87
Max	0.101	2.82	0.087	2.31	-0.016	-0.34	0.015	0.37
Average	-0.002	-0.09	-0.039	-1.07	-0.054	-1.22		
Panel size (min/max/av.)	5/8/6		5/8/6		1/2/1.6		1/2/1.6	
Error correlation within panels	0.7652		0.7994		0.7539		0.7567	
# obs	25302		30064		18461		20330	

Notes: Estimates are based on population averaged panel data probit model assuming equal error correlation within panels. z-values based on standard errors (robust conditionally on assumed correlation structure) for respective coefficients. [a] Marginal effects of explanatory variables on probability of positive outcome. Marginal effect for a dummy variable is calculated as increase in Pr(y=1) when respective variable changes from 0 to 1, while other variables (except those which are necessarily zero for the reference group) take their mean values. Reference groups not mentioned in the table: basic education or less; age 40-44; married or cohabited; no children; ethnic majority.

Source: Calculation based on LFS data.

Table 10 Determinants of labour force participation, controlling for studies and non-labour income

	Estonia				Lithuania			
	Men		Women		Men		Women	
Mean probability	0.694		0.577		0.671		0.579	
	dy/dx[a]	z	dy/dx[a]	z	dy/dx[a]	z	dy/dx[a]	z
Education: Higher	0.223	6.63	0.300	9.18	0.220	12.39	0.386	19.89
Postsecondary professional	0.196	3.84	0.245	6.44	0.192	10.77	0.268	13.66
Secondary general	0.148	5.37	0.171	6.28	0.118	6.53	0.193	10.39
Secondary vocational	0.152	5.02	0.229	7.31	0.189	10.56	0.260	11.61
Vocational	0.158	4.08	0.026	0.5	0.100	1.93	0.217	2.08
Age 15-19	-0.576	-10.3	-0.643	-10.98	-0.581	-17.74	-0.647	-20.00
(Age 15-19) × (Parents' wage per core family member)[b]	-0.066	-0.59	-0.126	-1.14	0.019	0.36	-0.064	-1.34
Age 20-24	0.004	0.12	-0.302	-7.54	0.014	-2.76	-0.092	-4.06
(Age 20-24) × (Parents' wage per core family member)[b]	0.079	1.24	0.037	0.61	-0.032	-1.41	0.024	0.41
Student/pupil	-0.251	-4.83	-0.321	-5.77	-0.307	-8.82	-0.245	-7.04
(Age 20-24) × Student/pupil	-0.356	-4.04	-0.127	-1.76	-0.336	-5.70	-0.367	-8.15
(Age 25+) × Student/pupil	-0.102	-0.92	0.118	1.52	0.129	1.60	0.071	1.41
Age 25 – 29	0.044	1.91	-0.173	-6.00	0.050	3.01	-0.028	-1.59
Age 30 – 34	0.027	1.02	-0.087	-3.42	0.040	2.41	-0.005	-0.34
Age 35 – 39	-0.012	-0.47	-0.029	-1.29	0.021	1.42	-0.019	-1.31
Age 45 – 49	-0.009	-0.30	-0.058	-2.19	-0.038	-2.3	-0.021	-1.31
Age 50 – 54	-0.009	-0.34	-0.118	-4.15	-0.068	-3.91	-0.063	-4.00
Age 55 – 59	-0.068	-2.21	-0.441	-11.07	-0.133	-6.82	-0.295	-14.11
Age 60 – 64	-0.480	-10.79	-0.716	-17.01	-0.463	-19.55	-0.657	-28.37
Age 65 – 74	-0.756	-16.93	-0.834	-22.00	-0.763	-30.54	-0.779	-36.14
Single	-0.042	-1.49	-0.007	-0.24	-0.147	-7.45	-0.044	-2.96
(Wage of spouse)/(family size)[b]	0.247	3.20	-0.049	-1.17	0.019	1.25	0.003	0.24
One child	0.110	3.77	-0.127	-3.71	0.110	2.24	-0.107	-5.65
More children	0.141	3.63	-0.307	-8.46	0.141	1.90	-0.173	-5.09
Ethnic minority	-0.052	-1.86	-0.097	-4.03	-0.034	-1.81	-0.073	-4.15
Disabled	-0.724	-14.65	-0.491	-9.12	n. a.	n. a.	n. a.	n. a.
Rural	-0.053	-2.43	-0.052	-2.63	0.050	3.46	0.003	-0.17
MW=Log (min wage) (last Q): main effect	-0.253	-1.33	-0.167	-1.17	0.324	2.77	0.625	5.16
(MW-mean(MW))×(age 15- 19)	0.682	1.33	0.363	0.81				
(MW-mean(MW))×(age 20-24)	0.296	0.51	0.709	1.41				
AW=Log (avg. local wage) last year: main effect	0.031	0.63	0.157	3.67	-0.053	-0.81	-0.110	-1.58
(AW-mean(AW))×age 15- 19							0.138	1.71
(AW-mean(AW))×age 20-24					0.186	2.76	0.246	3.00
(AW-mean(AW))×age 60-74						-1.69		
Log (last Q county unemployment rate)					-0.045	-1.97	0.032	1.35
County fixed effects	yes				yes		yes	
Panel size (min/max/av.)	1/2/1.7		1/2/1.7		1/2/1.6		1/2/1.6	
Error correlation within panels	0.7452		0.7337		0.7444		0.7474	
# obs	7432		8848		18461		20330	

Notes: Estimates are based on population averaged panel data probit model assuming equal error correlation within panels. z-values based on standard errors (robust conditionally on assumed correlation structure) for respective coefficients. [a] Marginal effects of explanatory variables on probability of positive outcome. Marginal effect for a dummy variable is calculated as increase in Pr(y=1) when respective variable changes from 0 to 1, while other variables (except those which are necessarily zero for the reference group) take their mean values. Reference groups not mentioned in the table: basic education or less; age 40-44; married or cohabited; no children; ethnic majority. [b] Parents' wage and spouse/partner's wage per family member are measured as deviations from their mean values divided by national average net wage.

Source: Calculation based on LFS data.

Table 11 Determinants of discouragement among inactive population aged 15-74.

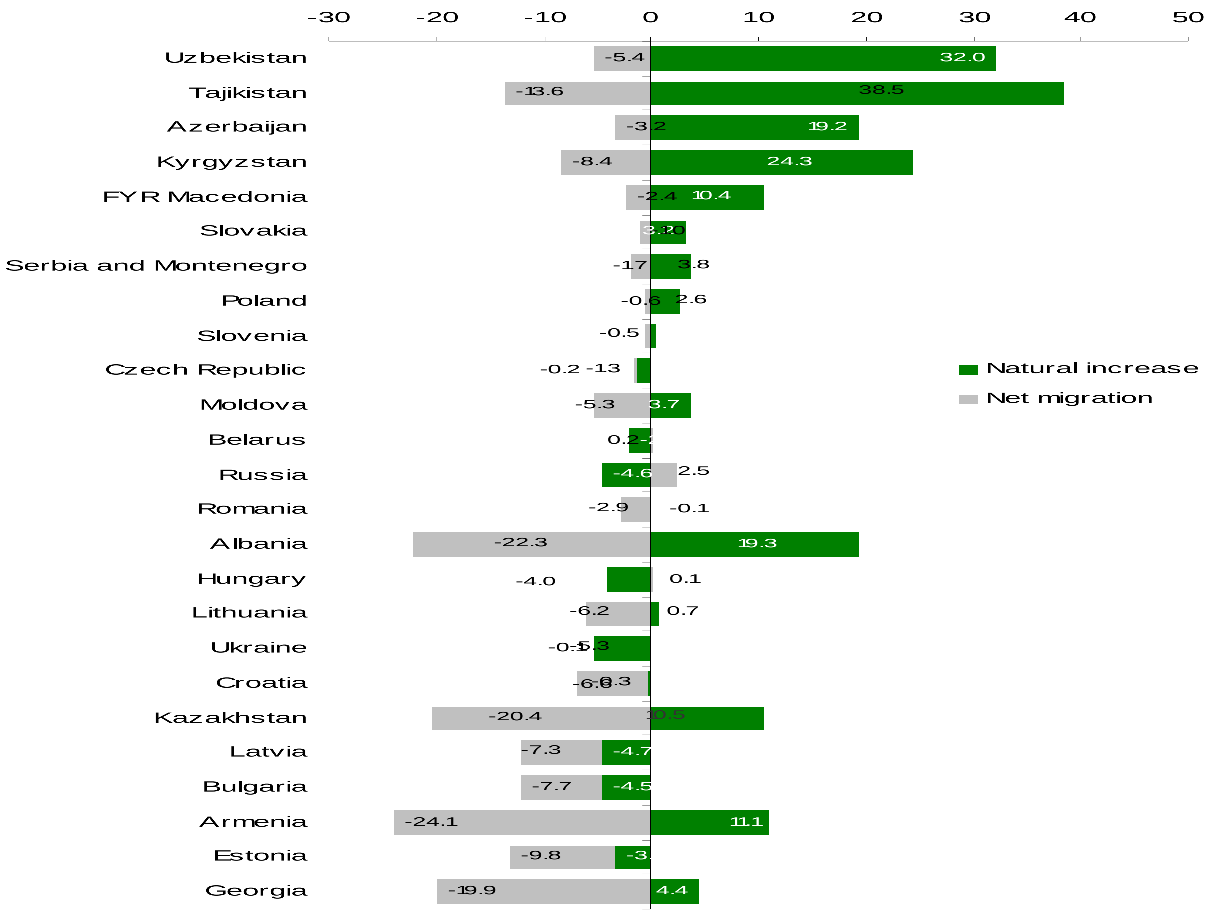
	Estonia, 2001				Lithuania, 2002-2003			
	Men		**Women**		**Men**		**Women**	
Mean probability	y=0.173		y=0.129		y= 0.049		y = 0.039	
	dy/dx[a]	z	dy/dx[a]	z	dy/dx[a]	z	dy/dx[a]	z
Education: Higher	-0.013	-0.33	-0.011	-0.49	0.013	0.88	0.012	1.41
Postsecondary professional	-0.021	-0.28	0.014	0.62	0.022	2.4	0.008	1.46
Secondary general	-0.015	-0.7	0.050	3.13	0.013	1.47	0.016	2.62
Secondary vocational	-0.018	-0.74	0.052	2.47	0.021	2.22	0.019	2.38
Vocational	0.012	0.36	0.056	1.91	0.046	2.09	-0.014	-7.93
Age	0.046	11.6	0.030	11.48	0.003	1.98	0.006	8.48
Age squared/100	-0.001	-11.57	-0.036	-12.44	-0.005	-3.27	-0.008	-8.86
Pupil/student	-0.081	-4.07	-0.036	-1.52	-0.024	-4.03	-0.013	-2.81
Single	0.011	0.54	0.019	1.07	0.000	0.04	0.004	0.94
One child	0.099	2.26	-0.022	-1.07	0.004	0.24	0.013	1.98
More children	-0.027	-0.56	-0.045	-2.60	0.006	0.20	0.009	1.42
Ethnic minority					0.000	-0.02	-0.001	-0.12
State language skills (vs. native speakers)								
Speaks	-0.044	-1.29	0.014	0.54	n. a.	n. a.	n. a.	n. a.
Understands, doesn't speak	0.087	1.45	0.032	1.18	n. a.	n. a.	n. a.	n. a.
Doesn't understand	0.066	1.87	-0.008	-0.41	n. a.	n. a.	n. a.	n. a.
Disabled	-0.134	-8.89	-0.065	-4.5	n. a.	n. a.	n. a.	n. a.
Rural (vs. cities except capital)	0.020	0.97	0.015	1.06	-0.003	-1.01	-0.005	-1.64
MW=Log (min wage) (last Q): main effect					-0.051	-1.01	-0.120	-2.78
AW=Log (avg. local wage) last year: main effect					-0.065	-3.66	-0.028	-1.79
(AW-mean(AW))×(age 15- 19)					-0.006	-2.47		
(AW-mean(AW))×(age 20-24)					-0.002	-0.82		
Log (last year county unemployment rate)					0.003	0.35	0.005	0.72
Age of max Pr(y=1)	41		42		31		41	
County fixed effects	yes		yes		yes		yes	
Panel size (min/max/av.)	1/2/1.7		1/2/1.7		1/2/1.6		1/2/1.6	
Error correlation within panels	0.3073		0.2887		0.3437		0.2604	
# obs	2513		3987		6366		8609	

Notes: The relaxed definition of discouragement applies: all persons who are willing to work and are available for work in two weeks time, but who are actively seeking job, are categorised as discouraged. Estimates are based on population averaged panel data probit model assuming equal error correlation within panels. z-values based on standard errors (robust conditionally on assumed correlation structure) for respective coefficients.

[a] Marginal effects of explanatory variables on probability of positive outcome. Marginal effect for a dummy variable is calculated as increase in Pr(y=1) when respective variable changes from 0 to 1, while other variables (except those which are necesserily zero for the reference group) take their mean values. Reference groups not mentioned in the table: basic education or less; married or cohabited; no children; ethnic majority.

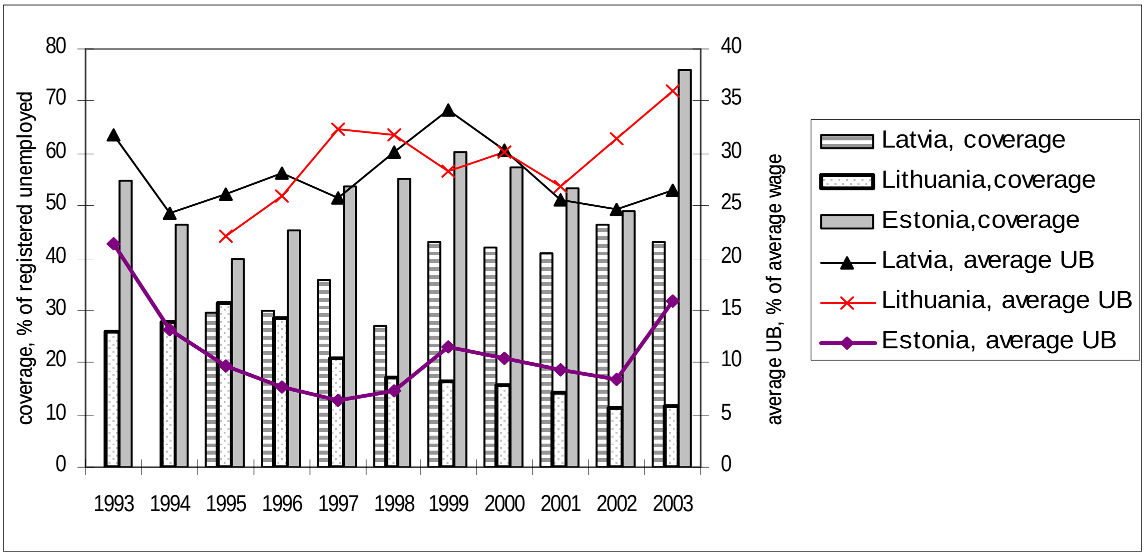
Source: Calculation based on LFS data.

Figure 1: Net Migration and Natural Increase by Country in the CEE-CIS Region, 1989-2002 (percent change)



Source: Heleniak (2004).

Figure 2. Unemployment Benefits in the Baltic countries, 1993-2003



Notes: Coverage is percentage of registered unemployed receiving unemployment benefits (UB). *Average UB* (after tax if taxed) is expressed as percentage of average net wage. The year 2003 point for Estonia includes both unemployment assistance benefits (UAB, flat at 7.8 percent of average net wage, coverage 52%) and new unemployment insurance benefits (UIB, coverage 24%, estimated average after-tax level 33 percent of average net wage). Maximal duration: 9 months for UB in Latvia and UAB in Estonia, 6 months for UIB in Estonia and UB in Lithuania. *Sources:* Estonian Labour Market Board (2004), Kuddo et al (2002), Statistical Office of Estonia (2004), State Social Insurance Agency of Latvia (2004), Central Statistical Bureau of Latvia (2004), National Labour Exchange of Lithuania (2004), Statistical Department of Lithuania (2004), own calculation.

Figure 3. Evolution of population, labour force, employment, and real GDP in the Baltic countries, 1989-2003

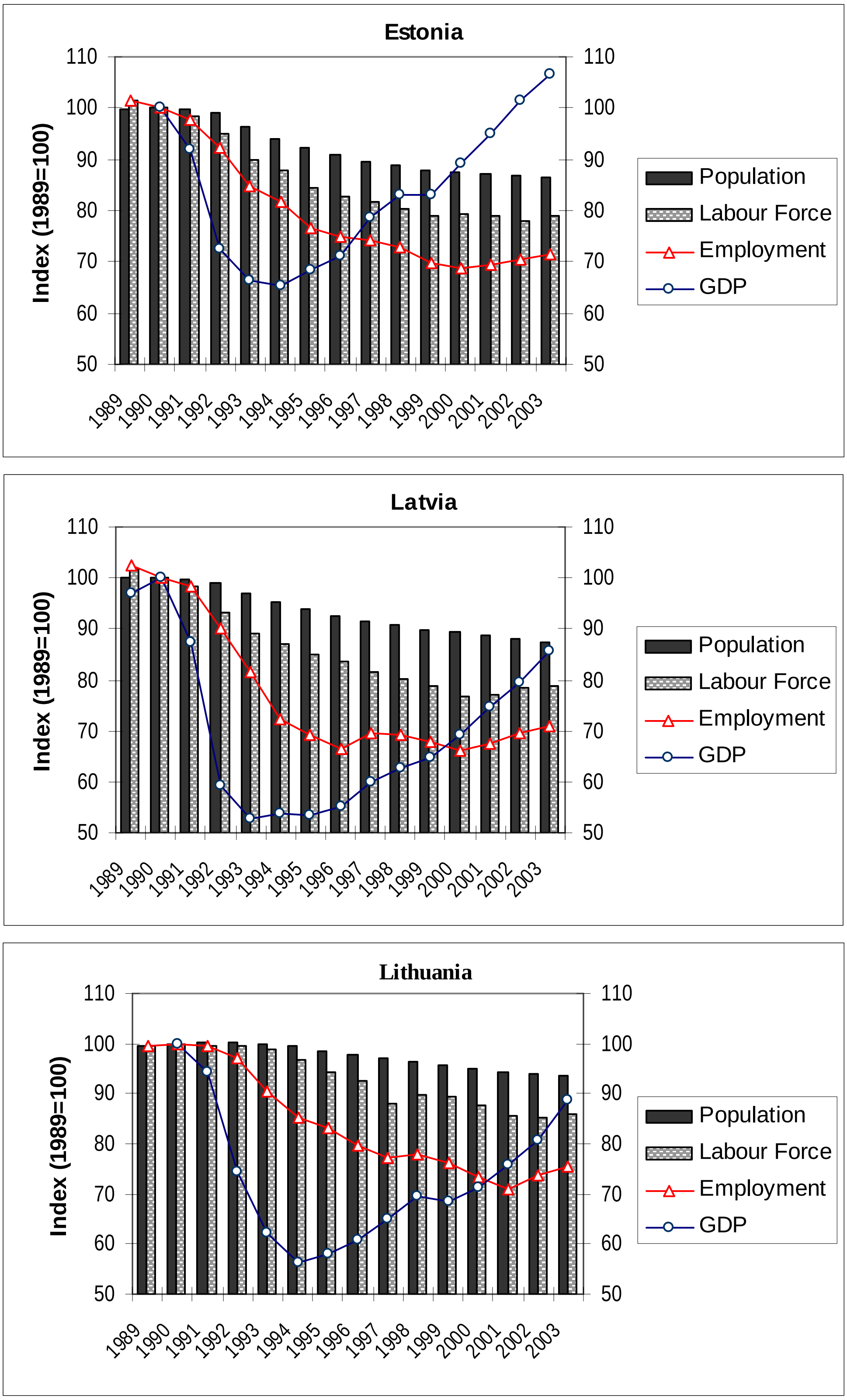


Figure 4. Labour market dynamics in Estonia, 1989-2003 (thousand population)

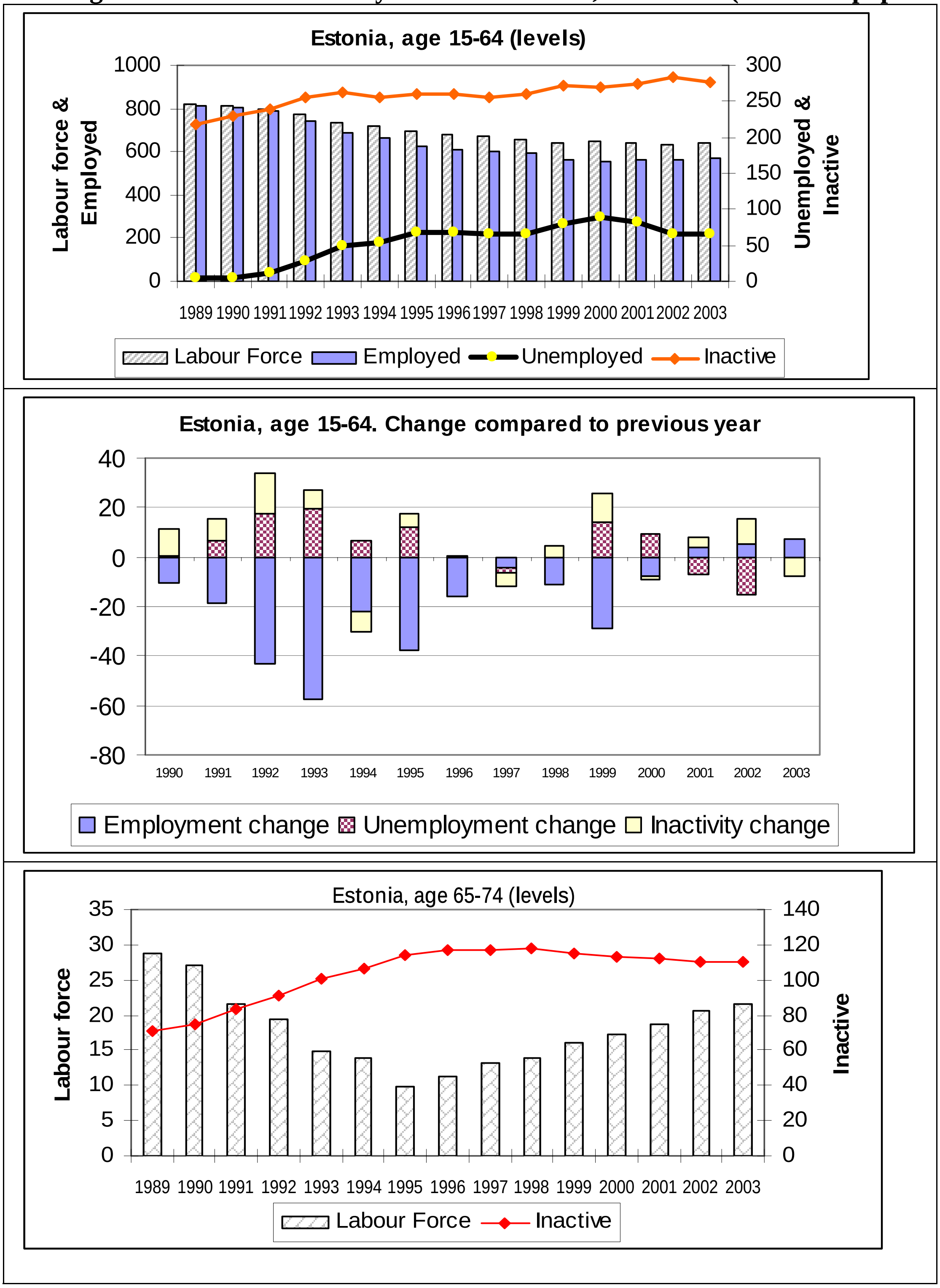
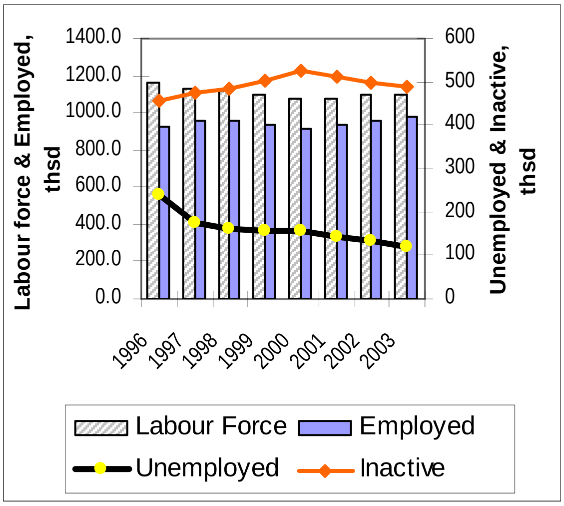
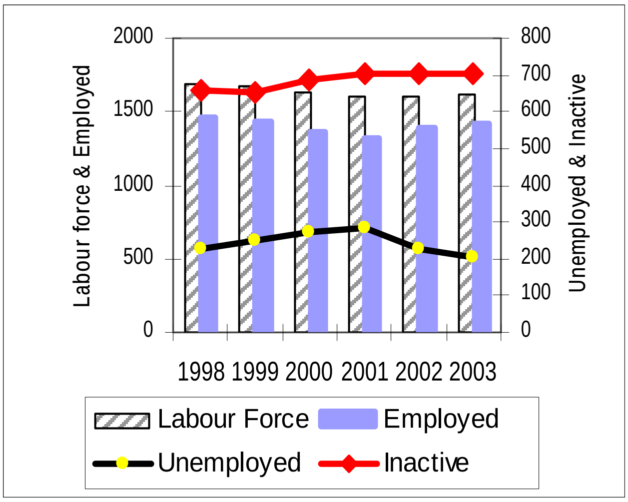


Figure 5. Labour market dynamics in Latvia (1996-2003) and Lithuania (1998-2003)

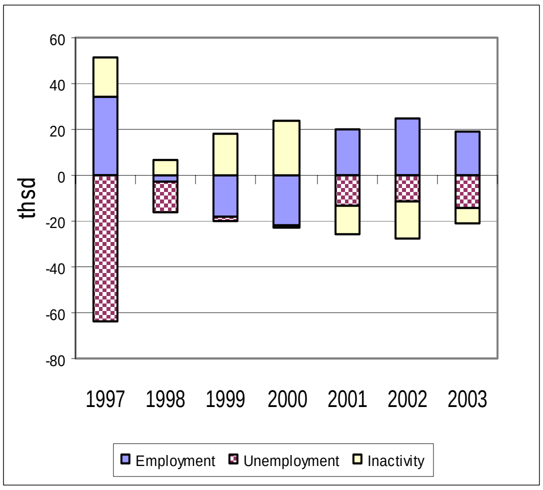
Latvia	Lithuania

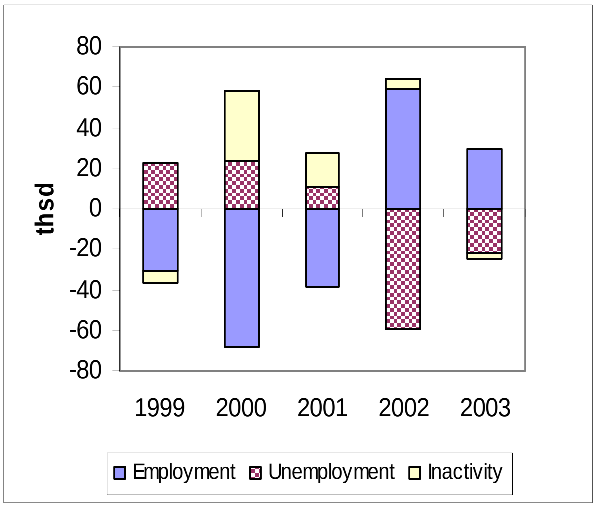
Age 15-64, levels



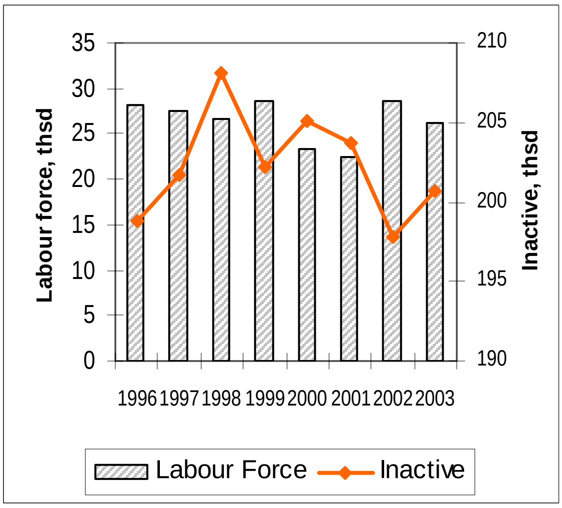


Age 15-64, change compared to the previous year





Age 65-74, levels	Age 65+, levels



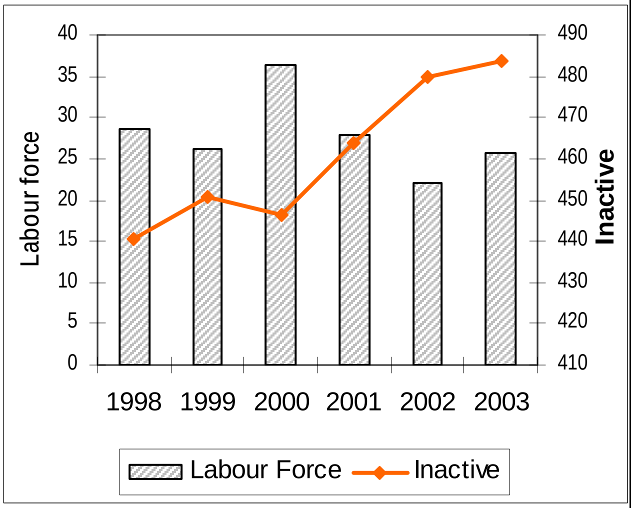
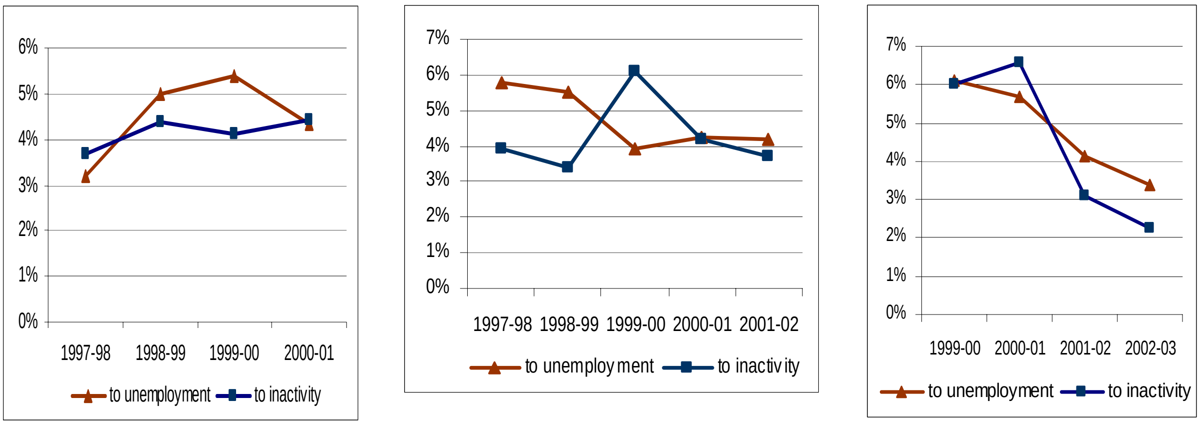


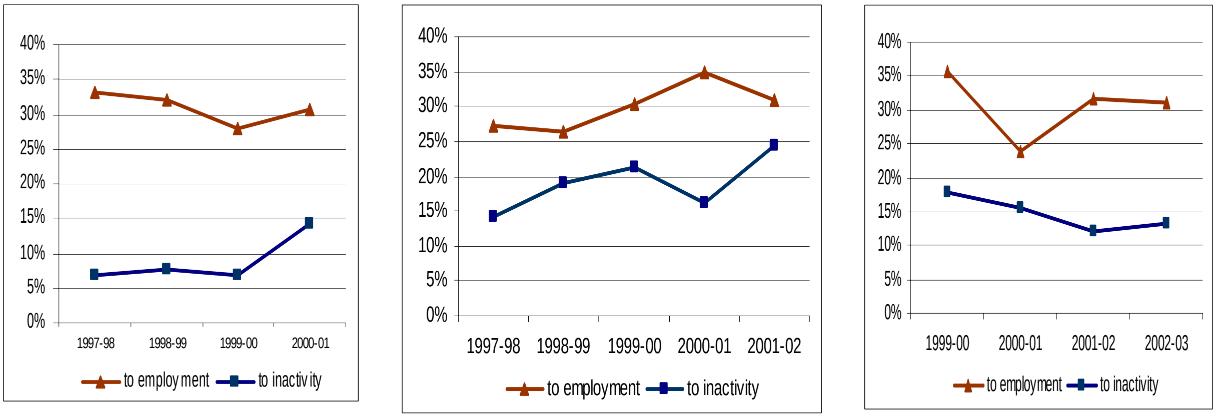
Figure 6. Labour market flows in Estonia (1997-2001), Latvia (1997-2002) and Lithuania (1999-2003

Estonia	Latvia	Lithuania

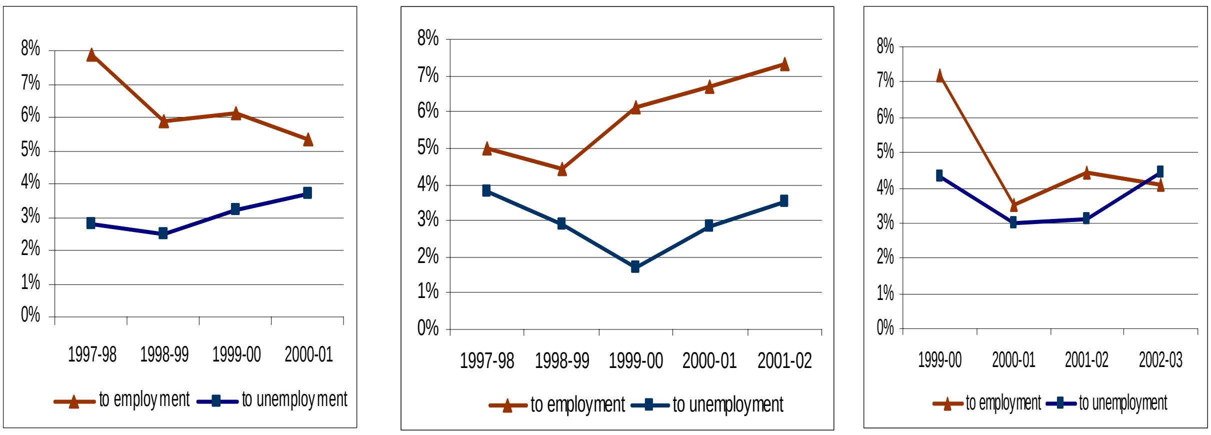
Outflows from employment by destination



Outflows from unemployment by destination



Outflows from inactivity by destination



Notes: 1997-2000 flows for Estonia are between Jan. of corresponding years, 1997-2001 flows for Latvia, and 1999-2001 flows for Lithuania are between Mays of corresponding years. Calculations were based on common sub-samples of the two LFS. The more recent flows (Estonia 2000-2001, Latvia 2001-2002 – annual average; Lithuania, 2001-2002 and 2002-2003 – average of Q2 and Q4) are based on the retrospective questions of the LFS. Estonia: population aged 15-74. Latvia and Lithuania: population aged 15 and older (for Latvian flows 2001-2002 only employed and unemployed aged 15-74 in 2002 were used, but since this group contributed 99.7% of employment and 100% of unemployment in 2001, results are comparable). Flows exclude the impact of migration, mortality, and new entrants who were younger than 15 in the first of the two periods. In this way, the impact of economic change is identified.
Sources: Calculation based on LFS data.

The Employment of the Roma – Evidence from Hungary

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Abstract. The paper is based on data of individual work histories of the 1993/94 representative Roma survey in Hungary. First the disappearance of full employment of Roma in the 1984-1994 period is documented by the use of a quasi cross-sectional macro model and the patterns of employment characteristics of the nineties are described. Then the erosion of employment is traced from individual histories controlling the effects of gender, age and schooling. Finally, particular aspects of low employment of Roma are accounted for, focusing on the role of low schooling, regional backwardness, and labour market discrimination.

Keywords: Economics of minorities and races, Discrimination, Regional inequalities, Transition
JEL Classification: J15, J7, R23

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1. Introduction

The economic transformation has put the greatest burden of all on the Romany population. As a result, the Roma have lost the basis of their living for the second time in the twentieth century. In the first half of the century, the disintegration of traditional Romany communities and the disappearance of the markets for traditional Romany crafts were both the products of a slow, evolutionary progress, which brought about – at least in part – an adjustment on the part of the Romany population in the long-run. As opposed to this process, the appearance of massive unemployment at the time of the transformation has wiped out in only a few years time just about all of the results of the slow modernisation. This modernisation had led to the integration of the Roma into Hungarian society – if only on its margins – through the expansion of the primary education and growth of industries based on uneducated labour. Undoubtedly, this integration was to a great extent only an illusion: the jobs offered by the distorted socialist modernisation could not last for long. Nevertheless, the social ascension of the Roma was real: a large number of people formerly on the margin of society were able to integrate into the society and have taken the first steps towards a more civilised life. In this process the spread of basic education was of crucial importance. The massive jobb loss of the Roma has made all of this history. With the collapse of the socialist economy, the market value of basic education has been nullified and a large part of the people that have integrated into society found themselves on the outside of society in a few years. The disappearance of surpassed forms of living, which happened at an unbelievably fast pace did not make it possible for the bulk of the Romany population to find successful forms of adaptation beyond bare subsistence. The more time the Roma spend in their current way of life, the stronger the vicious circle of poverty – low education – unemployment – poverty shall become. The situation of future generations is by no means more promising.

The Hungarian Roma are in a severe and unprecedented crisis. This paper was written in order to direct attention once again[1] to this acute crisis. This report is based on the employment histories part of the 1993/94 representative Roma Survey[2] which has been cleaned after many years of work, so we have at hand a previously unexplored database containing richer, more accurate information.

[1] Two earlier studies by the author focus on the employment of the Romany population based on a much narrower informational basis, see: Kertesi [1994], [1995].

The 1993/94 representative Roma Survey contained a bloc of questions on the employment histories of the adults in all of the households of the sample. We considered all those persons as adults who were at least 15 years of age at the time of the survey, and were not regular students in any educational institution. In the 2222 households questioned in the survey we had 5800 adults. Their employment histories represent the work histories of about 250-260 thousand Romany adults. These employment histories are made up of a chain of consecutive events from the person's first employment to his labour market status at the time of the survey at the turn of 1993/94. For all of those individuals who had never held a job in their life, this piece of information was recorded.[3]

The employment histories of the 5800 adults contained a maximum of 17 different spells, while an average employment history was made up of 3-4 spells. Our first aim was to make possible the comparability of these life stories differing in length – depending on the age and the type of employment history of the individual. The life stories of the 5800 adults were assembled of 21500 individual events, which contained the pieces of information given in detail in Footnote 3.

The cleaning of the database containing the employment histories took quite a long time. The correction of contradicting informations was in many cases only possible by checking the individual questionnaires, and we also went through the tedious work of checking the consistency of the employment chronologies with other pieces of information: the course of schooling, the changing of domiciles, the timing of births and so on. All of this work has successfully come to an end, and the database became adequate for statistical analysis.[4]

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[2] The survey, which comprises a 2 percent representative sample of the Hungarian Romany population was conducted by Kemény István, Havas Gábor and the present writer. For detailed information on the survey see: Kertesi–Kézdi [1998], chapters 1-3.

[3] For every event in the employment history of an individual we recorded the following pieces of information: 1. the starting and finishing year of the spell; 2. the type of activity in that spell (employment, unemployment, housewife, on child care leave, participation in education, member of the armed forces, in jail, retired); if the individual was employed: 3. what was her occupation; 4. how many months per year was she employed; 5. the industry of employment; 6. the settlement of the workplace; 7. the relation of the place of work to the place of habitation (same settlement, daily commuting, weekly, monthly commuting). We naturally had access to all the background information included in the Roma survey: the individual's gender, age, schooling, family status, the characteristics of the place of residence etc. So information on just about all the factors influencing labour market status were at hand.

[4] We formed three files from the original database, all representing different perspectives on the individual employment histories. 1. A file that contains a snapshot of the labour market status of the individual in the given year for all of the years from 1979 to 1994, this file shall be called "*snapshot file*". 2. A second file which measures the number of months employed, the length of the employment spells, the number of children born, the

The use of this database is particularly useful in reconstructing the dramatic crowding out of Romany workers from the labour market from the second half of the 1980s to the date of the survey, 1993/94. Our interest is not only motivated by the curiosity of the historian seeking to document the dissolution of a withered system – the disappearance of full employment – although this also is a not an unimportant goal. But the story is instructive to date: it helps understand *the structure and characteristics of Romany employment* which emerged from the ruins of full employment by the middle of the nineties.

The paper is organised as follows. First the disappearance of full employment of Roma in the 1984-1994 period is documented by the use of a quasi cross-sectional macro model and the patterns of employment characteristics of the nineties are described. Then the erosion of employment is traced from individual histories controlling the effects of gender, age and schooling, and particular aspects of low employment of Roma are accounted for, focusing on the role of low schooling, regional backwardness, and labour market discrimination. In the final section we summarise the basic findings.

fact of attending a night school in the period starting from the individual's entry unto the labour market to January of the given year (1979, 1980, …, 1994). This database shall be called "*flow file*". 3. In the third file our observations were not the individuals, but the events of their employment histories: this file contains altogether 21500 spells with all the relevant information about the events in the employment histories. Naturally, more than one spell (observation) of the same individual can be found in this file which we called "*event file*".

2. Romany employment between 1984 and 1993: a quasi cross sectional macro model

Consider the following two-state macro model (see *Graph 1*). The working-age persons in a given year (t) are in one of two labour market states: they are either employed or not employed. The increase in year (t+1) of the stock of employed in year t (E_t) can be attributed to two sources: those labour market entrants (mostly young), who have become employed in the given year (ye_t), and those from the stock of non-employed who have found a job in the given year (ne_t). The *total inflow into employment* is the sum of the two above flows: $ye_t + ne_t$. The stock of employed persons is reduced by two flows: those employed who have lost their jobs in the given year (en_t), and those employees who have retired (ep_t). By adding up these two flows, we receive the *total outflow from employment*: $en_t + ep_t$. Similar flows reduce and increase (in an inverse manner) the stock of non-employed (N_t).

Graph 1

Accordingly, the stock of employed (non-employed) in a given year (t+1) can be computed from the stock of employed (non-employed) in year t and the flows in year t by the use of the following equations:

(1) $$E_{t+1} = E_t + (ne_t + ye_t) - (en_t + ep_t)\,.$$

(2) $$N_{t+1} = N_t + (en_t + yn_t) - (ne_t + np_t)\,.$$

In the ideal case, information on the stocks can be found in cross-sectional databases, while the data on the flows comes from panel data. In our case, all of the information comes from the employment histories of the 5800 persons in the representative cross-section of the 1993/94 Roma Survey, so all of the data on the stocks in past years (E_t and N_t, where t = 1984,…, 1993) is taken from this database. Our estimates are based on the following procedure: we reproduced the transition matrices in *Table 1* for each pair of years from the "*snapshot*" file using frequency weights.[5]

[5] The Roma survey contains an about 2 percent sample of the whole Hungarian Romany population, so our frequency weights were of the order of about 50. The samples taken in Budapest and Miskolc are the exceptions because the sampling proportion in the first city was twice as large as in general, while in the second city it was four times as large. As a consequence the frequency weights used for the habitants of Budapest was about 25, while for the habitants of Miskolc it was about 12,5. The exact analytic weights differed from these values slightly due to the multistage sampling technique used. For further information see: Kertesi – Kézdi [1998] chapters 1 and 2.

Table 1

To make computations simpler, we considered all those as non-employed who were neither employed nor retired. In other words all unemployed persons, housewives, persons on childcare leave, in military service, jail and non-regular students were classified as non-employed. We note that the majority of the non-employed were unemployed, housewives or persons on childcare leave. Based on the transition matrices, we are able to estimate from our employment histories the stock of employed and non-employed persons as well as the labour market flows.

Our estimates are subject to some biases. E.g. stocks and the flows of the year 1984/85 do not contain those persons who have died since 1985, given that our information is based on the population of the year 1993/94. Due to this fact all of our estimates relating to absolute numbers are lower than the hypothetical estimates based on cross-sectional data. If we consider the biases of *relative* numbers, it is clear that the largest biases can be found in the estimates relating to the oldest cohorts, who are evidently made up of the retirees in a given year (ep_t and np_t). On the same grounds, it is easy to see that our estimates relating to the labour market entrants (ye_t and yn_t) are the least biased. Due to the number of deceased our estimates of the stock of employed and non-employed as well as the flows to and from these two states are biased to about the same extent, for the average ages of persons in these stocks are about the same in every year. Furthermore, those who are employed have on average more schooling than the non-employed, so we can expect the employed to have a lower mortality rate. In this manner, our estimates of the ratio of the stock of employed to the stock of non-employed should be considered as slightly upward biased in every year.

Graph 2

Graph 2 shows the time-path of the stock of employed (E_t), and non-employed (N_t). This graph makes clear the dramatic loss of Romany employment in the 1984 to 1993 period. In the middle of the eighties out of a working-age population of 160-180 thousand persons, there were about 120 thousand employed, and about 40-60 thousand non-employed. From the late eighties (1988-89) these proportions started to change gradually, so the stock of employed decreased first at a slow, then at a faster pace. By 1993 the stock of employed fell to about half (60 thousand persons) their number in the eighties, while the stock of non-employed (and not retired) grew by an enormous amount, to about 140 thousand persons. As a result, the employed/non-employed ratio, which was about 3:1 at the middle of the eighties, was worse than 1:2 in 1993.

We now decompose the change of employment relative to the stock of employed in the base year (in percentage) according to Equation (3):

$$(3) \qquad \frac{E_{t+1} - E_t}{E_t} = \frac{(ne_t + ye_t)}{E_t} - \frac{(en_t + ep_t)}{E_t}.$$

The first term on the right hand side (the *inflow rate*) stands for the pace of flow into the stock of employed in a given year *t*, while the second term (the *outflow rate*) stands for the pace of outflow from the stock of employed in the same year. Both these terms measure the percentage of growth (decrease) relative to the employment in the base year that can be attributed to the flow into (out of) employment. The time-path of the inflow and outflow rates can be seen in *Graph 3*.

Graph 3

Based on the evidence found in the inflow and outflow rates it is fair to say that the employment of the Roma was in a steady state at the middle of the eighties when low and stable in- and outflow rates maintained a relatively stable (and high) level of employment. This equilibrium destabilised at the end of the eighties: the outflow rate was about 7 percent in 1988 and this rose to 30 percent in four years (1992), while the inflow rate stood at 7-8 percent at the same time. As a result, the stock of employed decreased at an ever faster pace between 1988 and 1992.

In 1992 and 1993 we can observe the first signs of a new trend: the outflow rate ceased to increase, while the inflow rate doubled from 8 percent to 16 percent. In what follows we shall argue that – based on our fragmentary information – we can expect Romany employment to stabilise at the end of the nineties at a new (low-level) steady state. We can anticipate that this new steady state shall be characterised by in- and outflow rates about twice those of the steady state in the eighties, these rates will stabilise at around 15 percent. In other words, an employment pattern typical of the Third World could appear, where the level of employment of an uneducated group is not only very low, but the length of a typical employment spell is also very short and the stock of employed is alternating at a high speed. In this situation occasional work will be the dominant form of employment.

Graph 4

Now, we shall take a closer look at the components of the in- and outflow rates. *Graph 4* contains four panels: panel (*a*) shows the values of the outflow rate and its components – the rates of flow from employment to non-employment (en_t) and from employment to retirement (ep_t); while panel (*b*) shows the values of the inflow rate and its components – the flow from non-employment to employment (ne_t) and the flow of new entrants into employment (ye_t). In panel (*c*) we compare the flows between employment and non-employment (en_t and ne_t); while panel (*d*) concentrates on the rates of demographic change (ye_t and ep_t).

The changes in the structure of Romany employment are *basically* due to the changes in the rates of flow between employment and non-employment (see panel (*c*)), although the rates of demographic changes also altered somewhat in these ten years. This last development is not easy to see on panels (*a*) and (*b*), since the values of en_t and ne_t changed to such an extent between 1987 and 1993, that *in comparison* the changes in the rate of demographic change seem negligible. But panel (*d*) demonstrates that in the nineties the *balance* of demographic change is much lower than in the second half of the eighties: it fell from 2-4 percent to 1 percent. This difference can be attributed to both components of the demographic change: the rate of retirement in a given year (np_t) suddenly doubled after 1987/88 and stabilised at this higher level; while the employment rate of new labour market entrants deteriorated by one percentage point at the same time (it decreased from 6 to 5 percent and stabilised at that level).

The *net* in- and outflow rates that have been cleaned from the effect of demographic flows show the same time pattern as the *gross* rates. We distinguish three different periods in the decrease of Romany employment. In the first phase, between 1985 and 1989 a gradual erosion can be observed : the rate of flow out of employment (mostly job loss) steadily increased from year to year, it has risen from the level of 4%/year in 1984 to 7%/year by 1989 while the rate of inflow remained constant at around 3-4%/year. The second phase is the period between 1989 and 1992, when the pace of job loss increased by a staggering amount, from the 7%/year level in 1989 to the 25%/year in 1992, while the rate of inflow failed to increase. As a result, the decrease of employment – the balance of net in- and outflow rates – jumped from 3-4%/year to 20-21%/year. This last piece of information means that in 1992 the stock of Romany employed decreased by one fifth in only one year. We are only able to register the beginning of the third phase starting in 1992, when the rate of decrease of employment is easing. Although the rate of outflow has not stopped rising (from 25 to 38 %), the pace of this increase is slowing down. This phase is marked by the sudden jump in the inflow rate (from 4 to 11%). The net result of these two changes is the fact that in 1993 – for the first time since 1986 – the rate of decrease of employment is slower than the rate in the previous year. This may indicate that the market is beginning to approach a new steady state – at a very low level of employment (with around 50-60 thousand employed persons). The lack of data prevents us seeing at what exact value these flows will stabilise (if such stable state exists) in the second half of the nineties. Nevertheless the additional information on the *structure of employment* does suggest that after 1993/94 a new pattern of employment of Romany workers will emerge – characterised by *unstable employment and the dominance of occasional work.*

Graph 5

Look at *Graph 5*, where we measured the stability of Romany employment with the average length of an employment spell at the middle of the eighties and in the first part of the nineties. To describe each period, we chose three years and tried to answer the question: what was the typical length of the employment events in the individual histories in these two periods. The lengths of the employment spells were averaged over the three years and are measured in months per year.

We have to remark that the distribution of the length of employment spells in a given period is *independent* of the absolute level of employment in the given years. In principle, it is possible to have a situation where the level of employment is low – as it was in the first half of the nineties[6] – and at the same time most employment is secure (of 11-12 months per year length). In this case, the in- and outflow rates should be low, otherwise the representative spells of employment could not have been stable. According to an alternative scenario a low level of employment means at the same time a switch to occasional work of less than one year

[6] Based on the data in the year 1993: 60 thousand employed persons to a population of about 200 thousand working age persons not studying or retired means an employment rate of about 30 percent.

length. It is clear from *Graph 5* that the structure of Romany employment moved in this direction. The employment at the middle of the eighties meant the dominance of stable jobs – of 12 months/year length – while the employment of the first half of the nineties was made up of predominantly casual jobs of short duration. As opposed to the period of 1985 to 1987, when the ratio of long-run employment (12 months per year) was around 70 percent amongst Romany men, in the period of 1991 to 1993 the ratio of long-run employment fell to about half of that level (to 37-38%). A change of the same order came about in the structure of employment of Romany women.

This also means that in the middle of the nineties, the employment of Romany workers is not only characterised by its low level, but by the *high rate of in- and outflows*, so a pattern of *highly unstable* employment was in the making. Not only did the Romany population lose – once and for all – their jobs to a much larger extent than the average of the Hungarian population, and in this way were crowded out of the labour market, but those Romany persons who held on had to give up the hopes of a *long-term* employment relationship. The spread of unstable employment has caused social disintegration of those with a job: the lack of permanent employment also means the lack of a stable lifestyle, the continued presence of bread-and-butter worries, as well as a lower level of social transfers from the state and the employers – or even the loss of entitlements.

3. The collapse of full employment in a longitudinal perspective: Roma and non-Roma

Prior to their job loss during the economic transformation, the Romany workers driven out of the labour market – as *Table 2* shows – had long, continuous employment histories. Based on the evidence in *Table 2* we can say that the Romany workers who were crowded out of the labour market were not attached to the market to a lesser extent than those who were able to keep their jobs in the nineties. The length of continued employment spells before 1989 of the Roma still working in 1994 does not differ markedly from that of those out of work, neither among men nor among women, or in groups defined by age. Full employment meant about the same type of employment for Romany workers as for the rest of Hungarian society: stable, all year-long work. In other words: the dissolution of the full employment *started off from the same basis* for the Roma as for the rest of Hungarian society.

Table 2

The chance of job loss depends to a large extent on worker characteristics. With the collapse of the socialist economic model a large number of companies employing uneducated labour, manufacturing low-quality products and functioning inefficiently went bankrupt or contracted and the whole economy was forced into structural adjustments. The transitional crisis not only decreased overall labour demand, but it also altered the structure of demand: demand for low educated workers (with primary or vocational training school) underwent a dramatic decline, while the *relative* demand for labour with secondary (or higher) education increased. Furthermore: the employment crisis hit companies in the competitive sector much harder than the budgetary sector, so job loss was more frequent among blue-collar than among white-collar workers and in consequence struck the employment of men more than the employment of women.

The change in the structure of labour demand affected the Romany population particularly adversely because the typical Romany worker is blue-collar, of low schooling and male, just the type of person whose work has become the most devalued since the middle of the eighties. In comparison: the median Hungarian worker has finished secondary school, and has an equal chance of being male or female. Because of these differences the only way to correctly assess the disappearance of Romany employment is to do this in comparison to the employment of the typical Hungarian worker, with special attention to the *differing composition* of the two populations. To put it another way: we must control the most important attributes – gender, age and schooling – when accounting for the decrease of employment. This is what we shall do in this section.

Choosing a group of workers characterised by gender, age and schooling – for example the male workers with completed primary school and were 25-29 years of age in 1984 – *we follow the employment history of this particular group* from year to year in the period 1984 to 1994. Our question is: what percentage of the group would retain its employed status over these years. Naturally our chosen cohorts gain in age as time passes, so the men aged 25-29 in the above example would be 35-39 years of age by the end of our story in 1994. The passage of historical and of personal time (years of age) forces us to restrict our attention to those of 20-39 years of age in 1984, because they would be 30-49 years of age in 1994 and in this way would still be of working age.[7] As we showed in the previous section, this is the most important question: to what extent did the erosion of employment affect the *working age population*?

We chose 1984 as our starting point, because this probably was one of the "last years of peace" before the start of the transition in the labour market, so we can observe the "last stand" of full employment in the socialist economy. This is where a true long-run analysis should start off. We hope that it will be made clear in the discussion below that 1989 would not serve as a useful basis, for the gradual movement in the second half of the eighties foreshadowed the immense employment crisis after the economic transition (See Köllő [1998]).

We cannot grip the extent of job loss among Romany workers if we do not have a comparison group. This comparison will naturally be the whole of the Hungarian population. To our regret, a *longitudinal* database representative of the whole of the Hungarian population does not exist (nor a large sample of employment histories comparable to ours), which would make it possible to document starting from the middle of the eighties or from 1989 the impact of the

[7] In our study we mostly treat the group of persons aged 20-39 in 1984 at the aggregate level. We do not include the analysis of the employment of the five year birth cohorts, because their employment histories do not differ markedly. The case of the birth cohorts of all women is somewhat different for we found a gap of the order of 20 percentage points between the employment rates of the oldest and youngest cohorts in the second half of the

economic transformation based on individual employment histories. Because of this lack of data we shall have to be content with second-best methods. Nevertheless, our chosen method of analysis – that we do not analyse the employment histories of *individuals*, rather *birth cohorts* – made it possible to work out a second-best solution. If we take year-by-year large sample cross-section databases, and fix our analysis to cohorts, we have a quasi-panel database of these cohorts. This can only be done if we have representative large sample cross-sections for *almost all years*, so that the accidental random variations occurring because of differing sample designs can be smoothed by the *continuity* of the longitudinal database.[8]

Out of the 11 years of our period, we found adequate databases for 8 years (only years 1985, 1986 and 1988 are missing). Our sources of data were the following (in each case we had large individual files):

the 1984 CSO[9] Microcensus,
the 1987 CSO Household Expenditure Survey;
the 1989 CSO Household Expenditure Survey;
the 1990 CSO Census, 2 % representative file,
the 1991 CSO Household Expenditure Survey;
the 1992 CSO Labour Force Survey, simple average of the quarterly data;
the 1993 CSO Labour Force Survey, simple average of the quarterly data;
the 1994 CSO Labour Force Survey, simple average of the quarterly data;

In all the cases where we do not note otherwise, we calculated employment rates for the cohort aged 20-39 in 1984. In the following, we present our results by the use of *Graphs 6-11*. The use of graphs (as opposed to tables) is motivated by the fact that we simultaneously operate with four (sometimes five) dimensions: gender, age, schooling, ethnicity (Romany/full population) and historical time. *Graph 6* presents the path of employment by gender, indicating data for both Romany and full populations. We have the following observations.

Graph 6

1. In our ten-year period the job loss among Romany workers was even more dramatic than the (far from negligible) job loss in the whole population. As opposed to the middle of the

eighties, which is probably due to the timing of births. We found no such tendency on the graph of the Romany women. The graphs for the five year birth cohorts are available from the author upon request.

[8] Naturally, it is also important to form the schooling and labour market status categories in exactly the same way in all of the individual cross-sections.

[9] CSO = Central Statistical Office of Hungary.

eighties, when the employment rate of Romany male workers was not far from that of the whole population – it was only behind by 4-5 percentage points – a decade later this small difference grew to an enormous gap of 45 percentage points. A disadvantage of the same order accrued in the employment of Romany women by the middle of the nineties, although at the middle of the eighties Romany women aged 20-39 already had an employment rate 20 percentage points lower than all women. In ten years about two-thirds of the middle-aged Roma lost their jobs.

2. The rather moderate employment losses (of 10 percentage points) of the 20-39 year old women in the whole population was because a large proportion of women had a white-collar job in the budgetary sector, which was less hit by the transitional employment crisis. In contrast, Romany women were employed to a larger extent by the non-budget sector in blue-collar jobs, so they lost their jobs in about the same proportion as Romany men did.

Graphs 7,8

Graphs 7-8 show by gender the time path of the employment rate of the Roma and the whole population broken down by schooling categories relevant to Roma[10]: less than primary school, completed primary school, vocational training school[11]. The inclusion of the schooling variable makes it possible to draw a more detailed picture.

1.First of all we can say that the huge gap between the employment rates of Roma and the whole population is *not only due to differences of composition*. The situation at hand is not only because the Roma have much less schooling and as a consequence, they have lost their jobs to a greater extent. Although the graphs by schooling categories also show the effect of this difference in composition[12], the fact is that in 1994 in all but one[13] gender/schooling group Roma have a minimum10, and mainly a 20-30 percentage employment lag whereas this

[10] We left out of the analysis all those with secondary or higher education, because the number of observations in the Roma survey were too low to make detailed investigation possible.

[11] In the case of the Romany population we included all those persons aged 20-39 who had secondary or higher education as well as those with vocational training school to increase the sample size. For the whole population this category is comprised only of persons with vocational training school. This does not have any important effect on our findings. First, the number of Roma with secondary or higher education is minimal. Second, their inclusion in this category probably makes their employment situation look somewhat better than it actually is, but it still is much worse than the employment of the same category in the whole population.

[12] The differences between the graphs of the Roma and the whole population disaggregated by schooling category are much smaller than the gap between the graphs of the two populations not disaggregated by schooling (see Graph 6).

difference was in every case smaller than 10 percentage points ten years earlier. This fact makes clear that there are factors other than differences in schooling (as well as gender and age) which govern the differences in employment probabilities between Romany workers and the whole population. These factors can be of three origins: unmeasured characteristics affecting productivity, regional differences and discrimination in the labour market. We try to account for these factors in the next section.

2. An even more interesting observation in *Graphs 7-8* is that the employment gap is smallest in the totally uneducated category (less than completed primary school) – in comparison to those with completed primary school or vocational training school – whilst it is clear that Roma in this category are the hardest hit by regional backwardness. So the composition effect of the regional dispersion of the Romany population[14] plays a minor role in the widening employment rate differential between Roma and the whole population. This conjecture is confirmed by the calculations of *Section 3*, where we show that if the regional backwardness had the same effect on the employment probabilities of Romany workers as it does on the probabilities of the whole population with the same amount of schooling, then the regional dispersion of the Romany population would be of much less dramatic consequence on Romany employment than it is in reality.

In *Graphs 9-11* we include two further dimensions of our analysis: the employment rate of the *labour market entrants* and *early retirement*. The common characteristic of these two groups, – the market entrants and those of age potentially eligible for early retirement – is that both are *markedly exposed to the hazard of job loss*. Above we used the term "*potentially* eligible for early retirement" for all those working-age, but not young (over 35 years of age) persons, who are (1) severely ill or disabled; or (2) working in a job which if discontinued, does not accrue additional costs to the employer via side effects in other production-lines; or (3) in a marginal position in the internal or on the local labour market and do not have influential acquaintances in their community or workplace who would plead their cause. If the economy is in a crisis and jobs are destroyed, then it is least costly (and brings about the least conflict in the workplace) for employers to lay off these workers.[15] The situation is the same if it is not

[13] The only exception being females with less than completed primary school .
[14] To be more exact: the Romany population is over-represented in the village category and in the settlements with high unemployment regardless of the settlement size. The order of this over-representation is higher, the lower the schooling of the given Romany or non-Romany group.
[15] See the report of Fazekas and Köllő [1990] (pages 215-219.) on this phenomenon at the end of the eighties.

the employers who initiate the retirement, rather it is *the workers who seek refuge in early retirement from the menace of unemployment*. The laxity in the process of awarding disability retirements gives ground for this kind of behaviour. These laxities can go unnoticed on the part of the social security (or the state budget), because what is lost on disability pensions is saved on unemployment benefits, plus this way of dealing with workers without much hope of reemployment in the future does not put a burden on state-run (and provided by the local governments) welfare system.

The new labour market entrants are in a danger zone for similar reasons, and in particular those with neither high-level education nor uncommon professions. At a time of cut-back most companies also do not take on new workers. If the whole economy is in contraction, then the aggregate probability of employment of labour market entrants will decrease too. It is reasonable to expect that at the time of a crisis the chances of employment of market entrants will decrease *faster* than the chances of job loss of employees will increase, or even if the pace of change of these two probabilities would be the same, the entrants' chances of finding a job would start deteriorating *at an earlier date*. The reasons are similar to the case of early retirement: on the one hand it is less costly for the employer – ceteris paribus – to not hire somebody from outside than to fire a worker with some job-specific human capital, and he does not have to accrue the fixed costs of discharge; on the other hand the stop of hiring does not cause conflict on the inside of the workplace as opposed to firing.

This can be relevant here in two ways. First, there is in the Romany population a larger proportion of less healthy or less fortunate and of those in jobs easily dispensable than in the Hungarian population on average. Second, the Roma are less integrated into the local society or into the organisation of the workplace than the average person in Hungarian society. The consequence is clear in both cases: even if employers did not have preferences against Romany workers – simply because of working against weaker opposing forces – they would send them in greater proportion to early retirement or refuse hiring them. [16] (Naturally all of this is worsened by the discrimination against Romany workers in the marketplace.)

[16] No one has to think of some kind of a sinister plot. It is enough to consider the actual situation of admission or dismissal at a firm. In a case where there are no vacancies at a given firm, but there still are fresh graduates applying for a job, then an exception will only be made if a particular job applicant is supported by insiders (relations, friends working for the firm) or by outsiders having standing in the local society with connection to the firm's management. The same argument applies to lay-offs: those have a greater chance of survival, who have someone with authority standing up for them. These micro-scale decisions, which take place several thousand times shall have the consequence at macro level – without anyone's intention – that persons weakly

Graphs 9,10

Graphs 9-10 show the employment rates of labour market entrants[17], while *Graph 11* depicts the proportion of early retirees. We focused on the situation of entrants with completed primary school or vocational training school since they make up the bulk of the young Romany cohorts.[18] Our results show that it was fair to say that when overall employment is declining, the chances of employment for entrants are particularly bleak. We see the same phenomenon in both the graphs of Romany youths and of youths on average: the chances of hiring of labour market entrants decline to a larger extent – independently of their schooling – than does the probability of job loss of employees with the same level of schooling increase. Not only is the situation of entrants worse than the older workers', but the employment crisis affected them earlier. *Graph 10* depicting the relative situation of entrants belonging to the Romany and the whole population by schooling category also confirms our conjectures: the employment rate of Romany entrants starts to decline *at an earlier date* and – in particular for those with vocational training school – *to a greater extent* than their counterparts in the population as a whole.

Graph 11

The problem of early retirement is shown in *Graph 11*. This graph – just as the one depicting the situation of entrants – is based on simple *cross-sectional* data: it shows the percentage of persons already retired in the given cohort[19] in the given year (1984, 1989 and 1994). Because each of the cohorts is within working age, all of the data greater than zero is due to early retirement. There are three differing cases of early retirement: disability pensions, which

integrated into the society – like the Roma – shall have smaller chances of keeping or getting a job – all other factors held constant – than the average person in that society.

[17] We defined the category of new labour market entrants the following way. In the "snapshot file" of the Romany employment histories we considered entrants all those persons with completed primary education who had 15-19 years of age in the given year and whose starting date of their *first* employment history event was of the same year. For the whole population we could not register directly labour market entry from our cross-section files, we simply defined the date of entry by the use of birth date and years of schooling.

[18] We did not attempt analysing the situation of entrants with secondary of higher level education, because they represent a very small proportion of the Romany population, even in the youngest birth cohorts.

[19] In this case we *did not* follow the employment path of the given cohorts.

can be awarded to persons with a decrease of working capacity of the order of 67 percent[20]; early retirement, available to those within 3 years of retirement age whose employer provides pension payments until the age of retirement rather than firing the person; preferential retirement at reduced age, which is available to those working in jobs particularly detrimental to health and in some other professions (for example workers of the armed forces). Although we are not able to differentiate retirees by their source of entitlement, it is well known from aggregate statistics that persons on disability pensions make up the majority of early retirees. As already pointed out, the institutional sytem of the social security was a partner for a long time in supplying with this type of benefit workers who – with the loss of their jobs – had no other stable source of income. Applying for disability benefits has become one of the typical forms of *escape from unemployment*.

The fact that this way of escaping from unemployment was often used in the ten years between 1984 and 1994 is well documented in *Graph 11*. In 1994 – see panels (*b*) and (*d*) – the proportion of early retirees was almost the double of the proportion in 1984 amongst the men in the oldest three cohorts and was more than its double amongst women. It is highly unlikely that in these ten years the health of the Hungarian population has decayed at a pace that would explain the growth in early retirement. It is more than probable that this increase is due to job loss, which is a fairly well-known development of the economic transformation.

The story of the Romany workers is even more striking. (i) Early retirement has reached incredible rates in the Romany population. Although at the middle of the eighties the work histories of Romany workers already ended fairly often in early retirement – this is surely in connection with the health status of the Roma, who typically worked in jobs with unhealthy conditions and hard physical work – the fact that the proportion of early retirees in the five years between 1984 and 1989 among men aged 45-49 increased from 14 percent to 30 percent, among men aged 50-54 from 23 percent to 48 percent and among women age 45-49 from 13 percent to 30 percent indicates that in the case of Roma somewhat older than middle age early retirement was one of the *dominant* forms of job loss. (ii) The other characteristic of the Romany population is that the sudden increase in the proportion of early retirees

[20] In special cases persons with a decrease of working capacity of 50 percent were able to obtain disability pensions, but these persons were only entitled to a pension of much smaller value (the so-called temporary social allowance).

happened five years earlier than in the whole population, in the 1984-89 period.[21] It is particularly important to emphasise the *timing* of the flow of masses of Romany workers into disability retirement status, because in the same period according to the official Hungarian statistics there was hardly any unemployment and although there were lay-offs – mostly in jobs with low qualifications –, the proportion of these lay-offs was negligible. If we consider early retirement as a form of job loss – it makes no matter that it is the workers who apply for disability pensions – then we can say that the crowding out of Romany workers from the marketplace was fully in swing in the second half of the eighties, at the time of so-called full employment.

Finally, to end the discussion of the problem of labour market entrants and early retirement we have to underline that our evidence is in line with the observations we made based on the macro model in *Section 1*. In panel (*d*) of *Graph 4* the rate of inflow of entrants has started to decrease as early as 1986 – the order of decrease was 20 percentage points – and two years later (in 1988) with the increase of flow into early retirement the rate of outflow from employment doubled in just four years time.

[21] There was no significant change by 1994 compared to the data for 1989.

4. Accounting for the low employment: low schooling, regional backwardness and discrimination

In this section we take a look at the *consequences* of the developments of this crucial decade. Based on individual level cross-section data, we try to measure the role of low schooling, regional backwardness and discrimination on the stabilisation of the low employment rate in the Romany population. We used as a reference group the data of the September - October - November wave of the 1993 CSO Labour Force Survey, which contained – in this single wave - the additional question of ethnic origin. We excluded all those families from our sample, who were indicated as Romany by the interviewer, this way our reference group is representative of the non-Romany population of the country. Both the sample of Romany and non-Romany populations were restricted to persons of working age in 1993 – men aged 15-59 and women aged 15-54 – and we also excluded students of regular educational institutions. We considered all those employed in the Roma sample who worked as employees or as entrepreneurs and were not registered as unemployed in the year of the survey; for the non-Romany population, the category of employed was made up of persons who worked at least one hour in the week prior to the date of the survey and usually worked at least 10 hours per week plus were not registered as unemployed.

Low schooling

Tables 3-4 show the basic facts about the differences in schooling composition of the Romany and non-Romany working-age population and about the employment rates by schooling (plus gender and age) categories. *Table 3* shows the differences in schooling composition of the Romany and non-Romany population broken down by gender. It is clear that the Roma have much less schooling than the non-Roma – which is well known - but the magnitude of these differences is astonishing. Only 20 percent of Romany men have more than completed primary schooling opposed to the 65 percent in the reference group. This difference is even greater among women, with 60 versus 10 percent not having more than 8 years of schooling. This can be the cause of large differences in employment rates in itself. But it is made clear in *Table 4* that *schooling composition alone* – or even combined with gender and age – cannot explain the enormous gap in employment rates. Within almost all schooling categories – even after controlling for gender and age – we find differences of 20-30 percentage points in employment rates. There must be other factors than schooling at work here.

Tables 3,4

Regional backwardness

Another source of the disadvantage of Romany workers in finding employment might be the unfavourable regional dispersion of the Romany population. This might be due to two factors: Roma are over-represented in villages where the absence of work is more acute than in any other settlement category; and Roma are over-represented in those regions where employment is especially scarce – regardless of the type of settlement. The regional differences in the employment situation are well represented by the distribution of unemployment rates in the 170 labour office districts. In 1993 (when the national representative survey on Romany population was conducted) one can observe very large differences – of twenty to thirty percentage points in magnitude – between the unemployment rates of the micro-regions of the country.[22]

Tables 5,6

The regional disadvantage of the Romany population is documented in *Tables 5-8*. The difference in the geographic distribution of the Romany and non-Romany population broken down by settlement type in *Table 5*, by the rate of local unemployment in *Table 6*, by settlement type and unemployment rate combined in *Table 7*. Finally, we have calculated the raw differences in employment rates of the two populations by regions, that is by settlement type and local unemployment rate, which is given in *Table 8*.

Tables 7,8

The evidence in these tables clearly shows that the geographic distribution of the Romany population is extremely unfortunate from the viewpoint of employment possibilities. 60 percent of the adult Romany population live in villages (opposed to 35 of the non-Romany

[22] We calculated the unemployment rates for the 170 labour office districts of the OMK. The data used here is the unemployment rate for the third quarter of 1993. See Ábrahám – Kertesi [1998] for the exact calculations.

population), and both in towns and in villages – as well as in the country overall – they live in a considerably greater proportion than the non-Romany population in settlements severly hit by unemployment. The effect of this difference on employment possibilities cannot be overstated, as seen in *Table 8*. Both the employment probabilities of Romany and non-Romany workers are adversely effected by the local unemployment rate. It might well be that the local unemployment rate and the level of schooling of the population is in an inverse relationship and this amplifies the effect of the regional differences on employment. The fact is that the variance of employment probabilities across local unemployment rates is greatest within the village settlement type, where the differences in schooling are the smallest. This points to the importance of regional labour markets in determining the probability of employment, independent of the schooling level. The employment situation of the Roma is as bleak as it is, because *a large proportion of the Romany population live in regions characterised by deep economic crisis*.

Labour market discrimination

We refer to discrimination in those cases where the employers value workers of the same quality – with the same schooling, labour market experience and not differing in most other attributes (those of importance in their market productivity) - differently: they hire these workers with different probabilities or at different wages. There can be many kinds of causes to this discriminative labour market policy. According to the most accepted explanation the employers discriminate between individuals belonging to different groups because they believe, based on previous experience – be this belief well-founded or completely irrational – that in these groups they will find workers appropriate for their purposes with differing probabilities *keeping the workers' observable attributes fixed*. Evaluating a job applicant's expected productivity is a very difficult task, for it is a function of a number of not easily measurable individual characteristics[23] outside of the applicant's observable attributes. The appropriate selection at the *individual* level is all the more important the more schooling is needed for the particular occupation or the higher up the job is in the hierarchy. This is why employers not only ask for meeting few formal criteria from applicants to these kinds of jobs, but they try to come to know the applicant in detail (by the use of aptitude tests, persons or

[23] Next to cognitive abilities social skills like reliability, ability to co-operate, good-fellowship etc. also play an important part.

works of reference, in-depth interviews and the likes). This obviously is a very costly way of hiring personnel, which is not affordable in simple blue collar jobs with low qualifications.

If the employers try to make their decisions based on statistical regularities and expect to find acceptable workers in one group – e.g among the Roma – with lower probability, then they will use this group-level information in their decision, given that this is less costly than *screening at the individual level*. Most of the discrimination in the marketplace against Roma is of this – statistical – nature. It is not only a matter of the preferences of the employer for or against Roma - although this might also come into play for some individuals – when they decide about hiring a Romany worker, but it rather depends on *the relative cost of applying ethnic background as a screening device*. This makes the situation all the more difficult, for statistical discrimination leads to lower costs and in this way it is economically rational from the perspective of the employer – although it is morally and legally condemnable[24]. Even an employer without prejudice against Roma has to consider whether it is affordable to employ an expensive human resource management team if it is possible to screen applicants with a high reliability – although calculating with the costs of making wrong decisions sometimes – based on observable characteristics (like gender, age or ethnicity).

These kind of statistical judgements are *mixtures* of substantive observations and pure prejudice. It is nevertheless clear that there can be enormous differences in the aptitude, knowledge and skills of workers with the same schooling and experience. It is also clear that these differences have something to do with the schooling career of these individuals. For example those youths who finish primary school over-aged after several repeated years (and probably with bad results) will have on the average less (learned) skills, aptitude etc. than those who had a straight schooling career. If the schooling career of Roma is broken to a larger extent, then – given that this information is widely known – this gives grounds to prejudice against the whole group.

Table 9

[24] Not only is it condemnable morally, but legally too, for it is an inequity against the given person: even if it is true that persons in her group have a smaller probability of having some skill, she might be in command of the given ability – which is the condition of acceptance for the job - herself. The right to equal treatment requires that the process should treat her *as an individual*, not *as a member of some group*. By the same token it is clear why this law cannot be enforced easily: economic rationality and equitable human resources management are in conflict with each other.

It is a fact that Romany children stumble more often in their schooling career than the average child. If we take two randomly chosen persons with 8 years of education from the Romany and the non-Romany population then the Romany person has a much higher probability of having finished primary school over-aged, with repeated years and bad results. This is confirmed by the data in *Table 9*. (We only note in brackets – because it does not belong to the point of *this* study – that many of the broken schooling careers of Romany children can be attributed – at least in part – to some dysfunctional traits of the Hungarian educational system. The lack of primary schools or of resources in the small villages where the proportion of Roma is high; the growing segregation of Romany children inside the schools as well as the general incapability of the educational system to give adequate help to children with learning problems[25], all these factors contribute to the great number school failures among Romany children which in turn is one of the main causes of Romany unemployment.)

But if these statistical judgements do have real foundations, why do we still call this phenomenon discrimination? There are two reasons: first, because we should call discrimination all the cases where an individual gets treated according to the average expected characteristics of her group (and not her own characteristics), regardless of whether the statistical judgements about her group are "true". [26] Second, even if the differences attributed

[25] One extreme example of this dysfunction is that special schools – which can be considered as dead-ends of schooling careers – are filled to growing proportions by Romany children. For example, in Borsod county for the 1996/97 school year while the proportion of Romany children was around 17 percent in normal primary schools (own calculations based on Kertesi–Kézdi [1998], page 316.), then it was 90 percent in special schools (see: Loss–Páczelt–Szabó [1998]). These same proportions were 14.3 and 50.6 percent for the 1977/78 school year (Cigány tanulók [1978], pages 31. and 43.). The over-representation of Romany children in special schools grew from 3,5 times to 5,3 times in twenty years for this county.

[26] Even if the employer's practice is economically rational from his own point of view. A society can make the decision – by the way of her political representatives – to make the application of group level screening *more costly* – because it judges these morally inadmissible - through legal regulation and establishing institutions that guarantee the enforcement of rights. A sufficient law to counter discrimination would deter at least a part of the employers with powerful sanctions from the application of such practices. Although the Hungarian legal system is rather far from such a situation (not the sufficient laws, but rather institutions that guarantee the enforcement of rights are lacking), we can speak of hopeful first steps – these come only from non-governmental institutions. There is a method frequently applied in other countries of pointing out hidden discrimination (the *audit studies*), which has been first adopted in 1999 – in the case of the employment of a Romany person – by a legal aid bureau, the *Nemzeti és Etnikai Kisebbségi Jogvédő Iroda* (NEKI) (see: Fehér Füzet [1998] és [1999]). Given that this is a new and very important method, we take the freedom to present it briefly, based on Fehér Füzet [1998], pp. 12.: "The basis of this method – which is particularly useful in exposing problems in the labour and the housing market – is that a tester, who is a member of the given minority group and another one, who belongs to the majority, but otherwise has the same relevant [*observable*] skills and characteristics, pays a visit to the accused company or individual with the same goal, questions and requests. If the experience in this situation confirms the grievance – that is, the member of the minority group does not get the same reactions as her fellow majority tester, and the details of the testing procedure also attest that we have a discriminatory case at hand –, then we start off a legal procedure, where we use the documents of the audit and the testimony of the tester as evidence. " It is obvious that the consequences of a legal process like this are very important. On the experience of the audit studies see: Heckman–Siegelman [1992], Neumark [1996], and Goldin–Rouse [1997].

to these groups by the statistical judgements existed in reality, we cannot be sure of their effect on the future productivity in the job. To our best knowledge – probably because of the lack of data – there has been no attempt at measuring the effect of skills not captured by school attainment on labour market performance.[27] It is not clear whether at very low levels of schooling are there significant productivity differentials between individuals with successful and with unsuccessful schooling careers at all. But even if there are, we must point out: *no matter how small* these differences in expected productivity would be in reality, if they serve as bases to statistical judgements operated as a group level screening device, they would have the same effect on employment differentials *as if they were very large*. For the employment decision is made in a situation of uncertainty and it is a decision with binary choice (hire/do not hire).

It is clear from the discussion above, that no matter what method we choose to measure the extent of labour market discrimination, the measured effect will be a mixture of two components: the effect of unmeasured skills plus the "true" effect of discrimination. This is the consequence of the technique used to measure discrimination. The only way we can grip the differential valuation of labour of the same quality is to try to specify – to the best of our knowledge - all the individual and contextual factors having an effect on the probability of employment and in this way build a model within which we are able to control for the heterogeneity of the quality of labour. All of the phenomena that we cannot attribute to economic mechanisms, in other words all of the residual effects, we consider as the consequence of discrimination (or of the non-measurable elements of skills).

Table 10

The results of our attempt at measuring the effect of discrimination can be found in the equations estimating the probability of employment of *Table 10*. The equations contain the parameters of a host of individual variables (gender, age, schooling, family background: number of children and marital status) and a variable measuring the situation in the local labour market (the unemployment rate of the labour office district). We are interested in

[27] For such measurement very detailed data are needed, for example the results of ability tests which have been conducted before entry unto the labour market and earnings data for the same persons from several years later. In other words: a longitudinal database containing very fine data is needed. To our knowledge there are only a few these in the world. One of these is the database that has been used by an excellent recent study: Neal – Johnson [1996].

predicting the difference in employment probability between Romany and non-Romany workers, using these independent variables. We shall state our predictions relative to our reference category, that is those unmarried men aged 30-39, with completed primary school, without children, who live in a district with low unemployment rate (under 10 percent).

We base our predictions of the employment probabilities on the following calculations. Let us denote the vector of independent variables (1, $\mathbf{x}_1$, $\mathbf{x}_2$, x_3, $\mathbf{x}_4$, x_5, x_6) [28], where the variables are in turn: constant, four schooling dummies, four unemployment dummies, gender, five age category dummies, marital status, number of children. Let us denote the vector of estimated parameters $(\hat{b}_0, \hat{\mathbf{b}}_1, \hat{\mathbf{b}}_2, \hat{b}_3, \hat{\mathbf{b}}_4, \hat{b}_5, \hat{b}_6)$. Our reference category shall be fixed at men $(\hat{b}_3 = 0)$ aged 30-39 $(\hat{b}_4^4 = 0)$, not married $(\hat{b}_5 = 0)$, with no children $(\hat{b}_6 = 0)$, our interest is in the predicted employment probabilities based on schooling (*i*) and local unemployment rate (*j*), given by the following equation:

(4) $$\hat{p}_{klm}^{ij} = \frac{1}{1 + \exp(-(\hat{b}_0^k + \hat{b}_{1i}^l x_{1i}^l + \hat{b}_{2j}^m x_{2j}^m))} \quad .$$

The indices *k,l,m* – which can take on three values: *r* (Romany), *n* (non-Romany), . (missing) – shall denote whether the parameters of constant (*k*), schooling (*l*) and unemployment rate (*m*) variables are fixed at the values from the equation for the Romany (*r*) or the non-Romany (*n*) population or it is fixed at the reference value (.).

Based on the different predictions $\hat{p}_{klm}^{ij}$ we can evaluate a number of experimental situations: we can look at the predictions of employment probabilities of Romany and non-Romany workers based on different assumptions. Prediction $\hat{p}_{nnr}^{ij}$ gives the employment probability of a Romany (*m*=*r*) man aged 30-39 years, who is not married and has no children with *i* schooling, who lives in a district with unemployment rate *j* if the elements of his stock of human capital measured by schooling attainment as well as the elements not measurable were evaluated at the same level by the market as the human capital of a non-Romany male. We make this assumption operational by predicting the employment probability of Roma using the parameters of the constant term and the schooling dummies from the non-Romany

[28] The indices of variables $\mathbf{x}_1$ are i = 1,…,5 (0-7 classes, 8 classes=base category.,…, higher education), and the indices of variables $\mathbf{x}_2$ are j = 1,…,5 (-10 %= base category, 10-15 %,…,25+ % local unemployment rate), the indices of variables $\mathbf{x}_4$ are r = 1,…,6 (15-19 years,…,30-39 years=base category,…,55-59 years of age).

equation (k, $l = n$) in formula (4) instead of using the parameters from the Romany employment equation.

Table 11

The results of the predictions using different assumptions are summarised in *Graph 12* and *Table 11*. In *Table 11* the outcomes of five different hypothetical situations are shown. In the first three scenarios (the first three lines of Table 11) we fixed the unemployment rate at its lowest level and measured the Romany/non-Romany differences dependent upon (1) differences in the parameters of the constant term, (2) differences in the parameters of the schooling dummies, (3) differences in the parameters of both the constant and the schooling variables. In lines 4*a*-4*b* we fixed the parameters of the constant term and the schooling dummies at their values from the non-Romany equation and measured the effect of unemployment rates on the across-group differences in employment probabilities. Finally in lines 5*a*-5*b* we measured the combined effect of differences in the values of the constant term, schooling and unemployment dummies on the employment probabilities of persons with completed primary school and vocational training school.

Graph 12

The results of our predictions are documented in graphical form in *Graph 12*. The four panels of the graph contain the predictions of lines *3*; *4a*, 4*b* and 5*a*, 5*b* of Table 11 in turn. Panel (*a*) graphs the effects of schooling and the constant term with the local unemployment rate fixed at its lowest level (less than 10 percent). The differences – depending on the level of schooling – are of 11-17 percentage points in magnitude for the three lowest schooling categories which account for 97 percent of the working age Romany population.[29] As we emphasised above, this difference is the sum of two effects: the differences in human capital *within* a given schooling category and the discrimination in the labour market. We are not able to separate these two effects, but – as it is made probable by the data in *Table 9* – the effect of differences in human capital within schooling categories is not negligible.

Based on the evidence of *Table 11* – see the third line – at the aggregate level Romany men aged 30-39 (and living in regions with the lowest unemployment) have a disadvantage of 27

[29] This difference is 8 percent at the level of secondary school, and 13 at the level of higher education.

points in employment rates due to this component compared to non-Romany men with the same attributes. About half of this difference is the composition effect, which exists because the composition of Romany and non-Romany populations by schooling categories differs markedly (see *Table 3*), the other half of the disadvantage is due to differences in the predicted probabilities by schooling categories (this is the parameter effect). [30] We surely can say that *the differences in the probability of employment by schooling categories is due only to a smaller extent to discrimination* and that this is - for the most part – the consequence both of low schooling and disadvantages in other, non-measurable, skills.

As for the effect of the local unemployment rate, the picture is rather different. At first look, the composition effect does not seem negligible – while almost one fourth of the Romany population lives in districts with extremely high unemployment rates (higher than 20 percent) and more than half of them live in districts with an unemployment rate higher than 15 percent, then the bulk of the non-Romany population (more than two-thirds) lives in districts with less than 15 percent unemployment[31] - but the burden of the economic crisis would be a lot less heavy on the Romany population had its negative effect on their employment probabilities been of the same size as the effect on the employment of the non-Romany population with the same schooling.

The reference group, as before, is composed of males aged 30-39 who are unmarried and have no children. In districts with low unemployment the disadvantage in employment probability of both Romany workers with completed primary school and with vocational training school is not too large (of 16-17 percentage points).[32] If we supposed - *following just one line of thought* – that *all* of this disadvantage of 16-17 percentage points is due to *differences in quality* - that is, to unobservable skills – still it is hard to explain *why this gap is growing with the worsening of local unemployment*. The fact that in districts with higher unemployment rates the relative employment probability of Romany workers is declining – see panels (*b*)-(*c*)-(*d*) of *Graph12* – is a sign that the crisis of the local economy hit the employment of the Romany population much harder than the employment of non-Romany people with the same gender, age, schooling, and family background. This difference is substantial: in the case of workers with completed primary school the gap grows from a base of 16 percentage points to

[30] The details of the methods of decomposition can be found in the Appendix.
[31] See Table 6.
[32] See Table 12. Panel (*a*).

32 percentage points in the districts with the highest unemployment (see panel (*c*)) while in the case of workers with vocational training school the difference grows from 17 to 40 percentage points (see panel (*d*)). *It is hard to interpret this phenomenon as a sign of anything other than discrimination in the labour market*. Our data bear witness to stronger discrimination in those parts of the country, where the competition for jobs needing only low qualifications is strong and the employment problems of majority workers with low schooling can be relieved at the expense of Romany workers searching for a job.[33]

[33] In these districts the proportion of the Romany population is higher than the average proportion for the whole country: see the maps in the Appendix of Ábrahám-Kertesi [1996]!

5. Conclusion

Based on individual employment histories, we tried to document the crowding out of Romany workers from employment in the ten years between 1984 and 1994. With the use of a quasi cross-sectional macro model, we demonstrated that the employment of working age Roma fell from 75 percent to 30 percent in ten years. We put forward the hypothesis that the employment of Romany workers at the middle of the nineties was not only at a very low level, but was characterised by high in- and outflow rates and an employment pattern – known from the Third World – with unstable employment and short employment spells was emerging. Not only did most of the Romany population lose their jobs to a much larger extent than the average of the Hungarian population, but those Romany persons who held on had to give up the hopes of a *long-term* employment relationship. The spread of unstable employment has caused social disintegration of those with a job: the lack of steady employment also means the lack of a stable lifestyle, the continued presence of bread-and-butter worries, as well as a lower level of social transfers from the state and the employers – or even the loss of entitlements.

We also traced the crowding out of Romany workers from the market along the individual employment histories, comparing this development to the situation of the non-Romany workers. We observed a growing gap between the employment possibilities of the two populations (to the disadvantage of the Roma), that cannot be fully attributed to the differences in the composition of the two populations. The Roma have lost their jobs to a far greater extent not only because they have much less schooling, but we suspect that along with their disadvantageous regional dispersion, discrimination in the market place against them also plays an important part. We pointed to a few regularities in the employment of new labour market entrants and early retirees suggesting the presence of discrimination. We also presented evidence that the job loss of Romany workers through early retirement had already started in the second half of the eighties, at the time of so called full employment. The labour market consequences of the economic crisis hit the Roma first, yet none of the companies or industries first swept out by the crisis had a particularly high proportion of Romany workers.

Finally, based on individual cross-sectional data, we tried to compare the relative weights of the different causes of low employment: low schooling, regional disadvantages and discrimination. With equations predicting the probability of being employed, we demonstrated that about half of the differences in employment probabilities depending on the type of schooling were caused by the effect of differences in the composition of the Romany and non-Romany populations by schooling. Our analysis of regional disadvantages pointed out that although the effect of differences in composition is sizeable, these disadvantages have a much more depressing effect on the employment of the Roma than on the employment of non-Romany workers with the same attributes. It would be hard not to interpret this phenomenon as a sign of discrimination. Based on our research we can say that the employment prospects – and from another viewpoint: life chances – of the Romany population are rendered feeble by basically three factors: low schooling, regional disadvantages and discrimination. All therapy should work to mitigate these forces. [34]

[34] See Kertesi [1995] and Kertesi-Kézdi [1996] for details of some earlier proposed policy reforms.

Appendix: The Oaxaca-Blinder decomposition of employment probabilities

Let us denote the distribution of the Romany (r), and the non-Romany (n) population by schooling (i) and local unemployment rate (j) f_r^{ij}, and f_n^{ij}. Naturally:

$$\sum_i f_r^{ij} = \sum_i f_n^{ij} = \sum_j f_r^{ij} = \sum_j f_n^{ij} = 1 .$$

We denote the predicted employment probabilities by $\hat{p}_{klm}^{ij}$, where i (i = 1,…,5) represents the given schooling dummy, and j (j = 1,…,5) the given dummy for the local unemployment rate; whereas k, l and m – which can only take two different values: r = Romany, n = non-Romany – tell us whether we fixed the parameters of the constant term (k), the schooling (l), and the unemployment rate (m) variables at the value taken from the equation for the Romany (r) or the non-Romany equation (n), or at value for the reference group (.) when making the employment probability predictions. The exact expression for the predicted probabilities is the following:

$$\hat{p}_{klm}^{ij} = \frac{1}{1+\exp(-(\hat{b}_0^k + \hat{b}_{1i}^l x_{1i}^l + \hat{b}_{2j}^m x_{2j}^m))} \quad \text{(where: } i, j = 1,\ldots,5 \text{ and } k, l, m = r, n, .)$$

For example, the prediction $\hat{p}_{nnr}^{ij}$ makes it possible to quantify what employment probability – depending on the local unemployment rate – a Romany man with i schooling, aged 30-39, not married and having no children if we used the constant term and the parameters for the schooling dummies taken from the non-Romany equation for making the prediction. (This means that we assume that the schooling and unmeasured skills of Romany men were valued at the same level on the market as the characteristics of non-Romany men.).

Now, using the above predicted employment probabilities, and the data on the distribution of the Romany and non-Romany populations by schooling and local unemployment rate, we are able to decompose the aggregate differences in employment probabilities depending upon schooling by the use of equations (1), (2) and (3), while the differences depending upon the local unemployment rate can be decomposed according to equations (4*a*) and (4*b*), or (5*a*) and (5*b*). In every case, the first component reflects the composition effect, while the second the parameter effect. We used two kinds of decompositions for every question: in the case of the first decomposition we used the non-Romany parameters for calculating the composition effect, while in the second we used the Romany parameters in the calculation.

1.
$$\begin{aligned}\sum_i f_n^{i1}\hat{p}_{nn.}^{i1} - \sum_i f_r^{i1}\hat{p}_{rn.}^{i1} &= \sum_i (f_n^{i1} - f_r^{i1})\hat{p}_{nn.}^{i1} + \sum_i (\hat{p}_{nn.}^{i1} - \hat{p}_{rn.}^{i1}) f_r^{i1} \\ &= \sum_i (f_n^{i1} - f_r^{i1})\hat{p}_{rn.}^{i1} + \sum_i (\hat{p}_{nn.}^{i1} - \hat{p}_{rn.}^{i1}) f_n^{i1} .\end{aligned}$$

2.
$$\begin{aligned}\sum_i f_n^{i1}\hat{p}_{nn.}^{i1} - \sum_i f_r^{i1}\hat{p}_{nr.}^{i1} &= \sum_i (f_n^{i1} - f_r^{i1})\hat{p}_{nn.}^{i1} + \sum_i (\hat{p}_{nn.}^{i1} - \hat{p}_{nr.}^{i1}) f_r^{i1} \\ &= \sum_i (f_n^{i1} - f_r^{i1})\hat{p}_{nr.}^{i1} + \sum_i (\hat{p}_{nn.}^{i1} - \hat{p}_{nr.}^{i1}) f_n^{i1} .\end{aligned}$$

3 $$\sum_i f_n^{i1}\hat{p}_{nn.}^{i1} - \sum_i f_r^{i1}\hat{p}_{rr.}^{i1} = \sum_i (f_n^{i1} - f_r^{i1})\hat{p}_{nn.}^{i1} + \sum_i (\hat{p}_{nn.}^{i1} - \hat{p}_{rr.}^{i1}) f_r^{i1}$$
$$= \sum_i (f_n^{i1} - f_r^{i1})\hat{p}_{rr.}^{i1} + \sum_i (\hat{p}_{nn.}^{i1} - \hat{p}_{rr.}^{i1}) f_n^{i1} .$$

4 *ab* $$\sum_j f_n^{ij}\hat{p}_{nnn}^{ij} - \sum_j f_r^{ij}\hat{p}_{nnr}^{ij} = \sum_j (f_n^{ij} - f_r^{ij})\hat{p}_{nnn}^{ij} + \sum_j (\hat{p}_{nnn}^{ij} - \hat{p}_{nnr}^{ij}) f_r^{ij}$$
$$= \sum_j (f_n^{ij} - f_r^{ij})\hat{p}_{nnr}^{ij} + \sum_j (\hat{p}_{nnn}^{ij} - \hat{p}_{nnr}^{ij}) f_n^{ij} .$$

5 *ab* $$\sum_j f_n^{ij}\hat{p}_{nnn}^{ij} - \sum_j f_c^{ij}\hat{p}_{rrr}^{ij} = \sum_j (f_n^{ij} - f_r^{ij})\hat{p}_{nnn}^{ij} + \sum_j (\hat{p}_{nnn}^{ij} - \hat{p}_{rrr}^{ij}) f_r^{ij}$$
$$= \sum_j (f_n^{ij} - f_r^{ij})\hat{p}_{rrr}^{ij} + \sum_j (\hat{p}_{nnn}^{ij} - \hat{p}_{rrr}^{ij}) f_n^{ij} .$$

In the case of decompositions 4*a* and 5*a* we calculate the distributions and the predicted probabilities for the group of persons with 8 years of schooling (i = 2), while in decompositions 4*b* and 5*b* the same is done for the group with completed vocational training school (i = 3).

We predicted the employment probabilities with the parameters taken from the logit equations in *Table 10*; the distributions were taken from the same data. When calculating the distributions we had to make the following simplifications to get around problems stemming from small cell size: in equations (1), (2) and (3) we calculated the schooling distributions for men aged 30-39 living in districts with an unemployment rate of less than 10 percent (which means that we did not disaggregate by marital status and number of children); while in equations (4*a*) and (5*a*) we calculated the distribution of men aged 30-39 with 8 years of schooling by local unemployment rate categories (once again we did not disaggregate by marital status and number of children); finally for equations (4*b*) and (5*b*) we calculated the same distribution for men with completed vocational training school (in this case, we were not able to disaggregate the sample either by age, or by marital status and number of children). These simplifications might bias our results somewhat – especially for equations (1), (2), (3), (4*a*) and (5*a*) – but we are convinced that the magnitude of these biases is ignorable.

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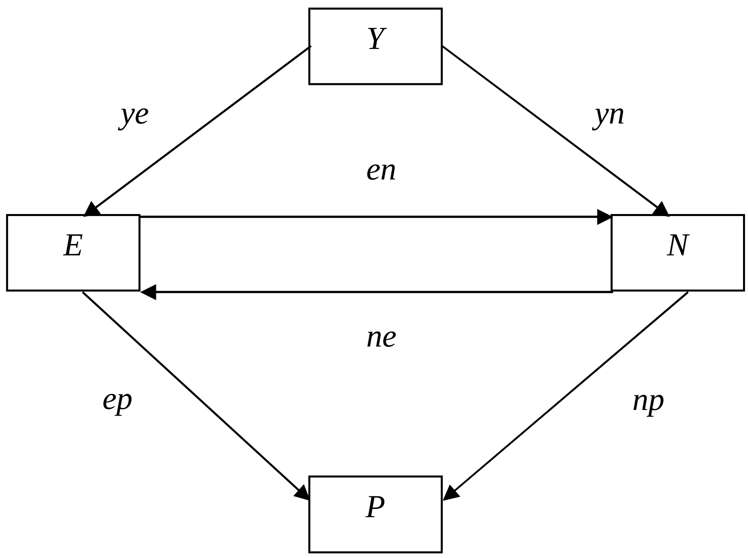
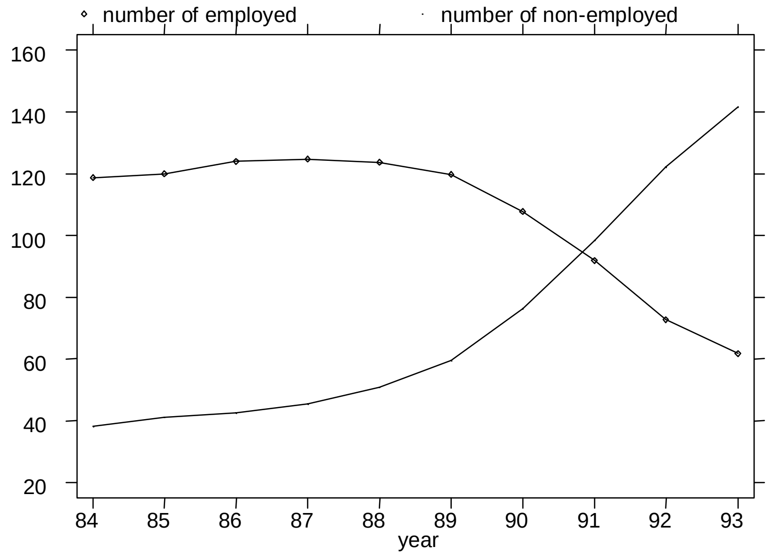
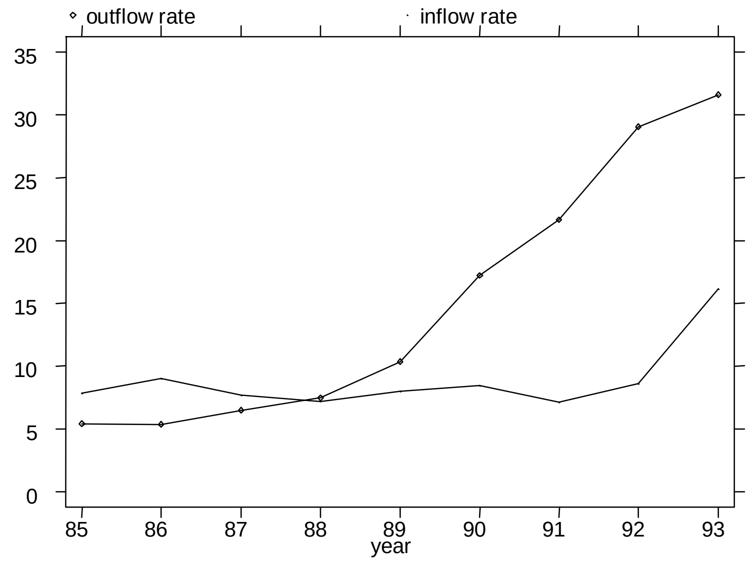
Graph 1: Labour market stocks and flows



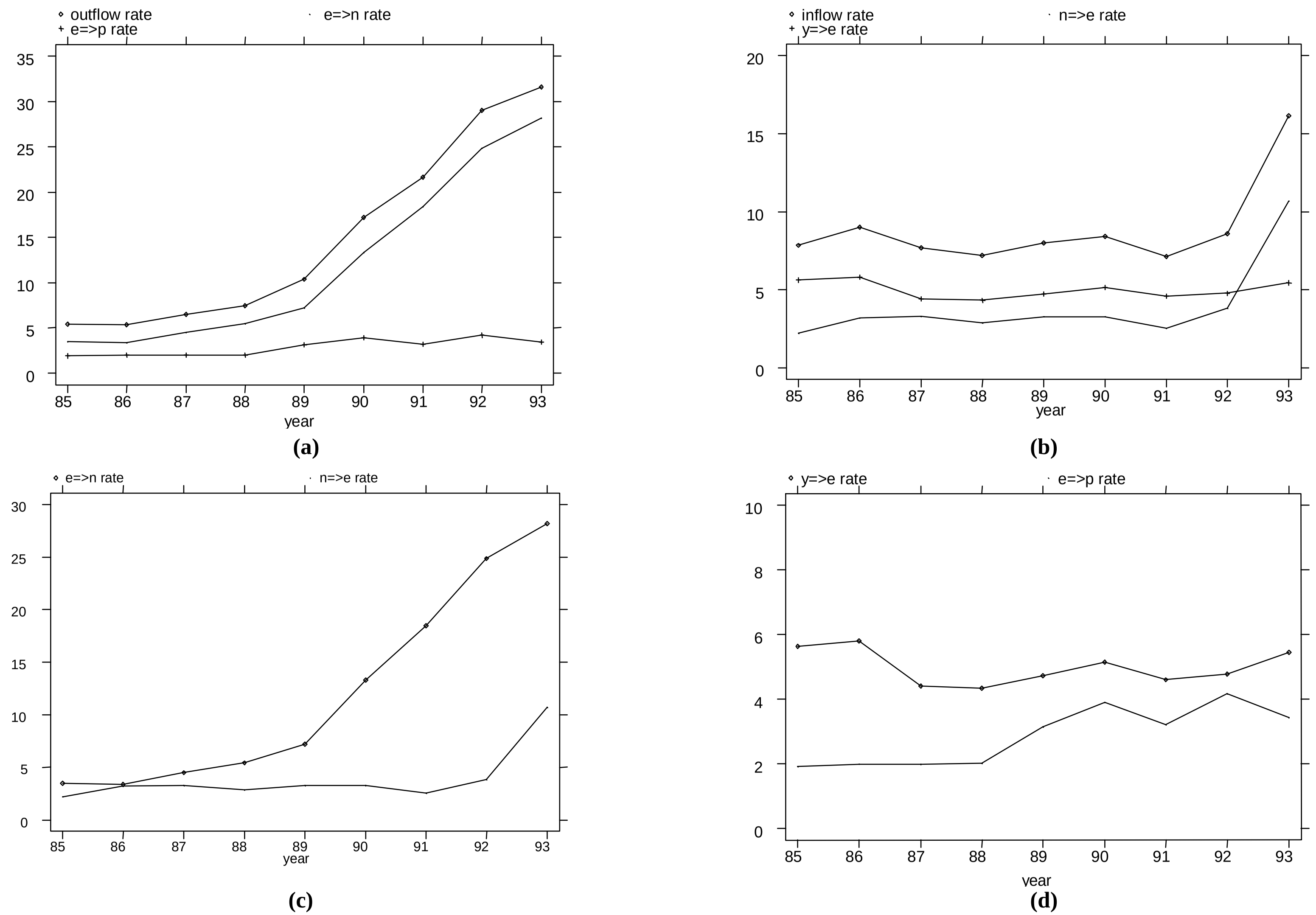
Notes: Y = labour market entrants
E = employed
N = non-employed
P = retired



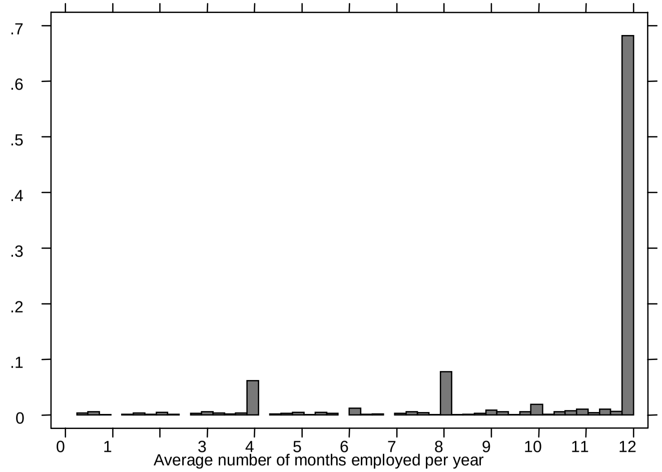
Graph 2: Changes in the number of employed and non-employed persons between 1984 and 1993 (thousands of people)



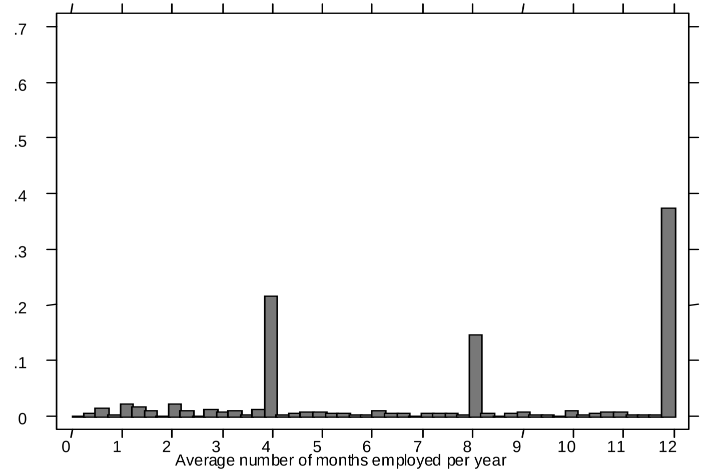
Graph 3: The yearly rates of flow into and out of the stock of employed between 1985 and 1993 (per cent)



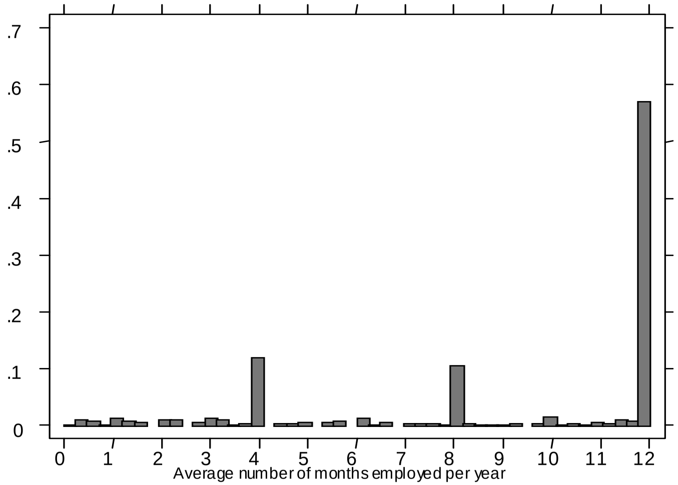
Graph 4: The yearly rates of flow into and out of the stock of employed persons between 1985 and 1993 (per cent) (e: employed; n: non-employed; y: new labour market entrant (young); p: retired)



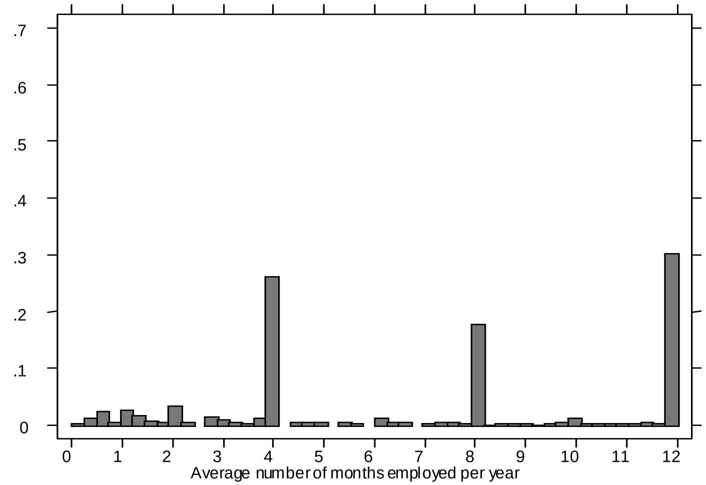
Men, 1985 – 1987



Men, 1991 – 1993

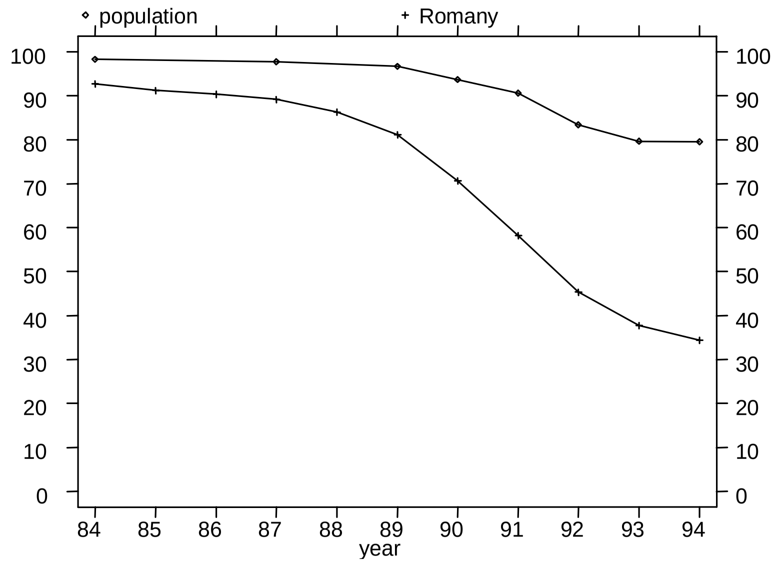


Women, 1985 – 1987

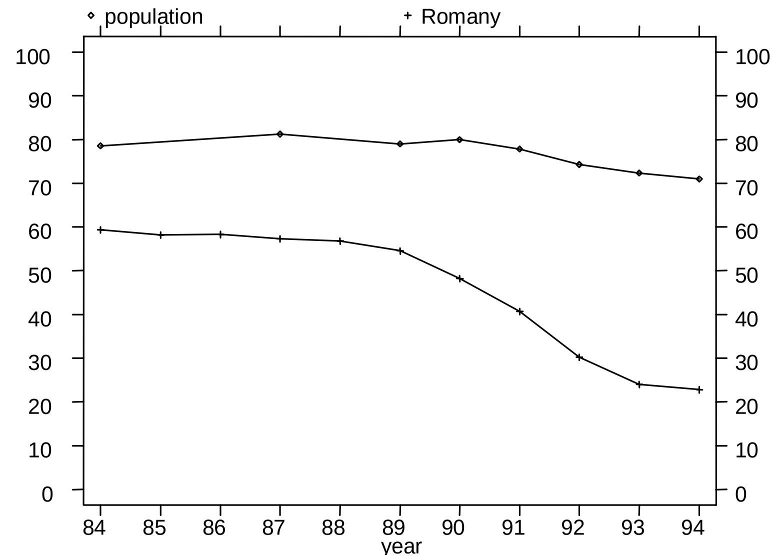


Women, 1991 – 1993

Graph 5: The average number of months worked per year by those who were employed in the period of 1985-1987 and 1991-1993:

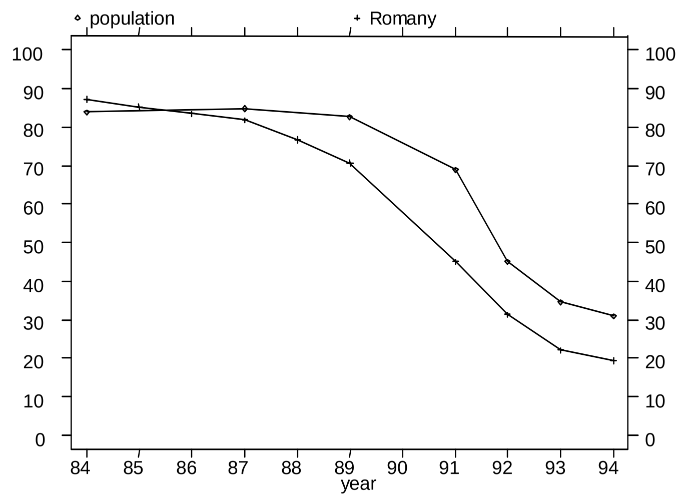


(a) men

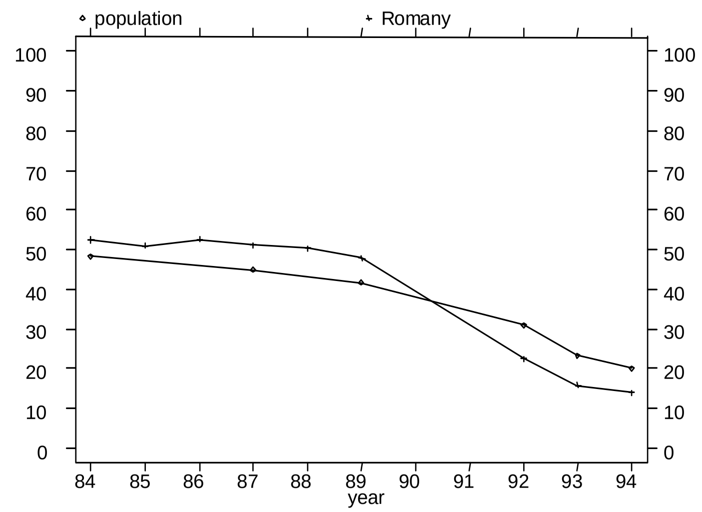


(b) women

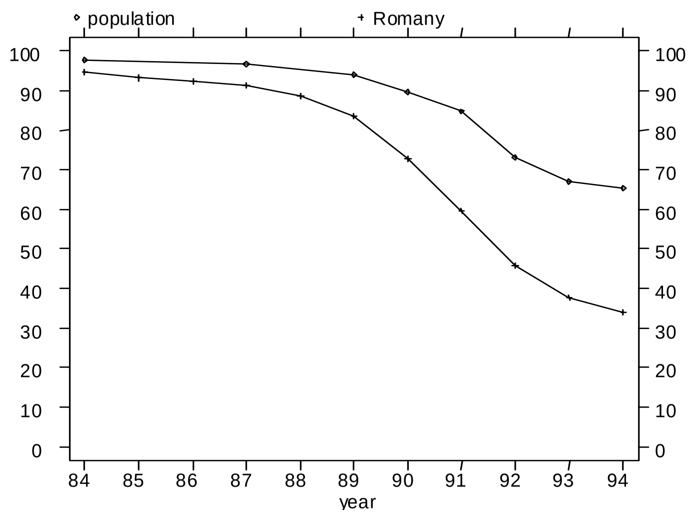
Graph 6: What percentage of the population aged 20-39 kept their jobs between 1984 and 1994 depending on ethnic origin and gender?



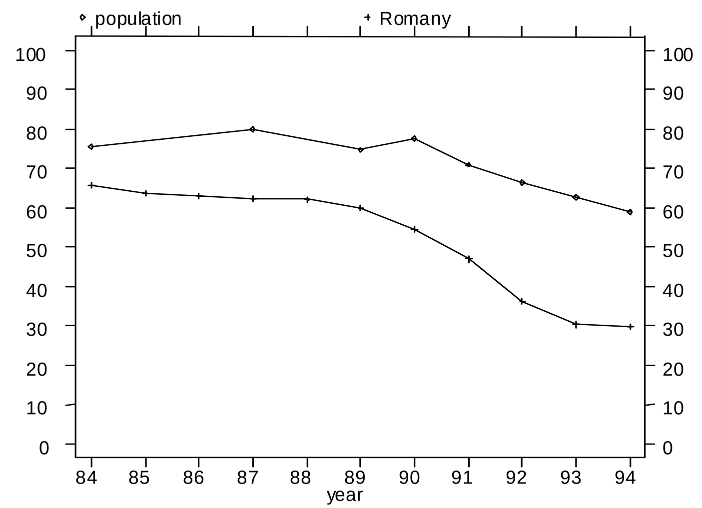
(a) men with 0-7 years of education



(b) women with 0-7 years of education

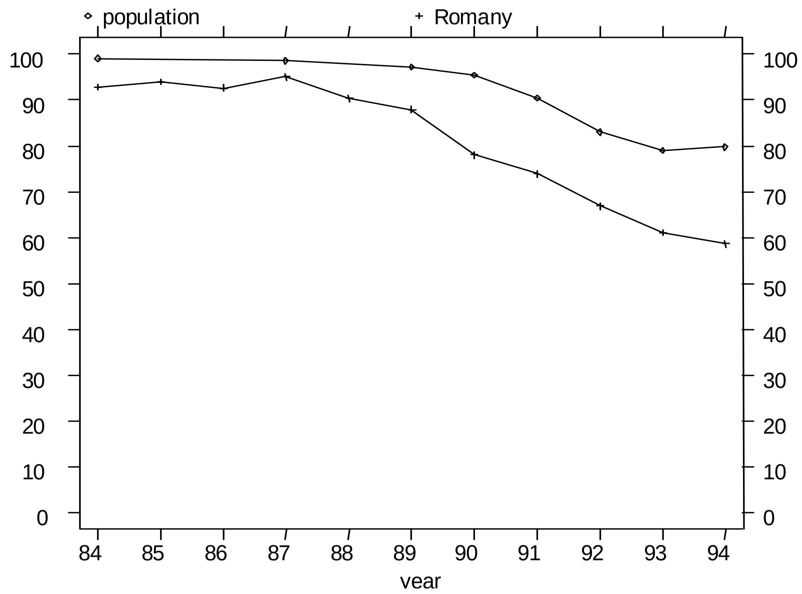


(c) men with 8 years of education

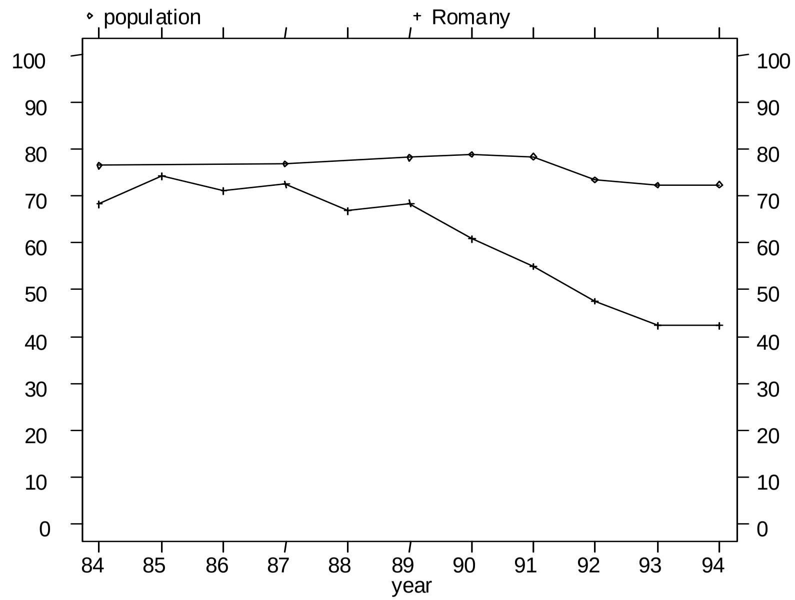


(d) women with 8 years of education

Graph 7: What percentage of the population aged 20-39 kept their jobs between 1984 and 1994 depending on schooling, ethnic origin and gender?

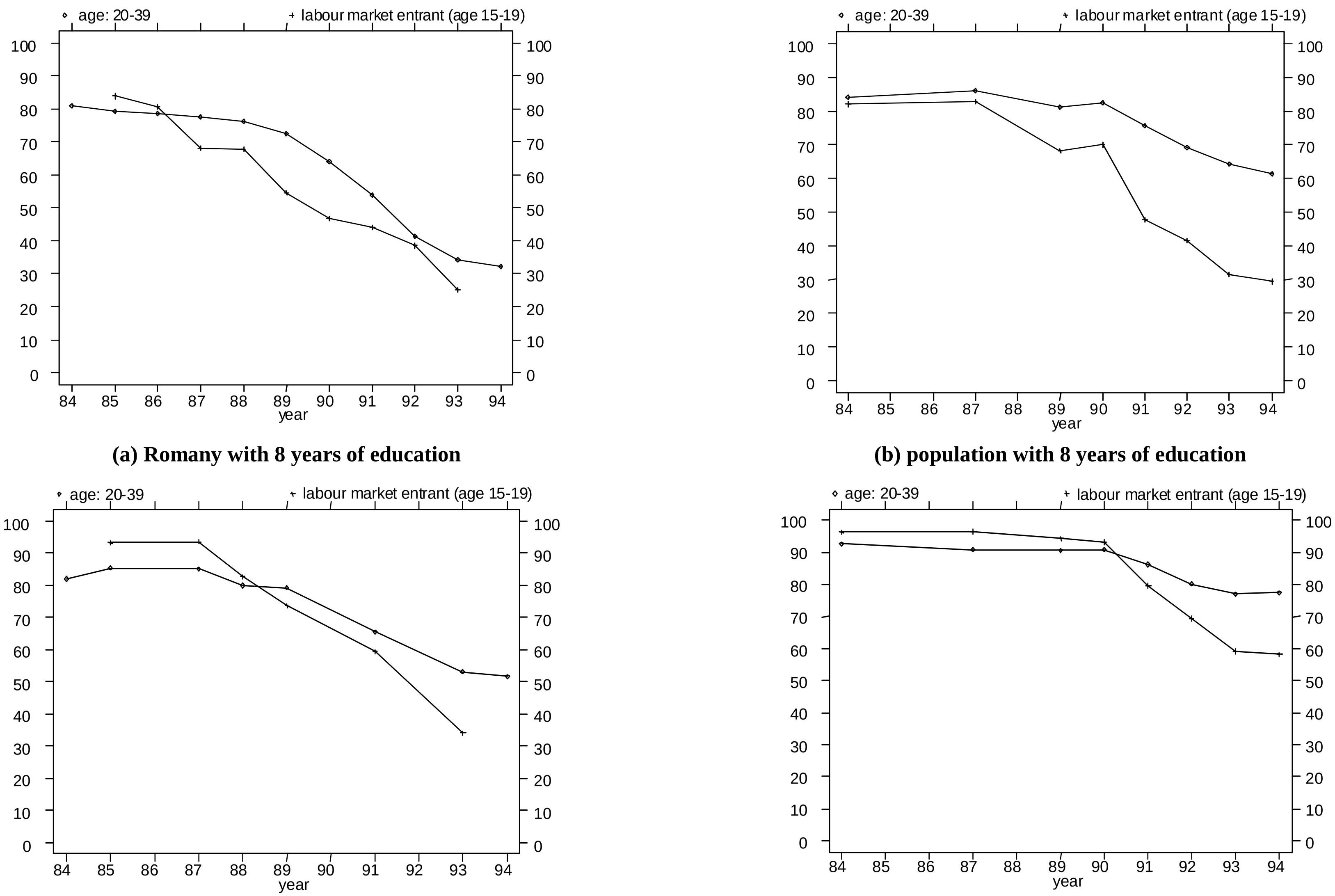


(a) men



(b) women

Graph 8: What percentage of the population aged 20-39 with completed vocational school kept their jobs between 1984 and 1994 depending on ethnic origin and gender?



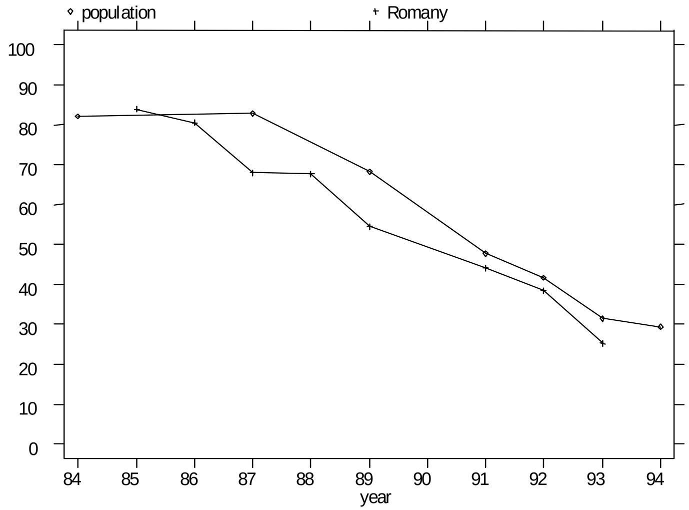
(a) Romany with 8 years of education

(b) population with 8 years of education

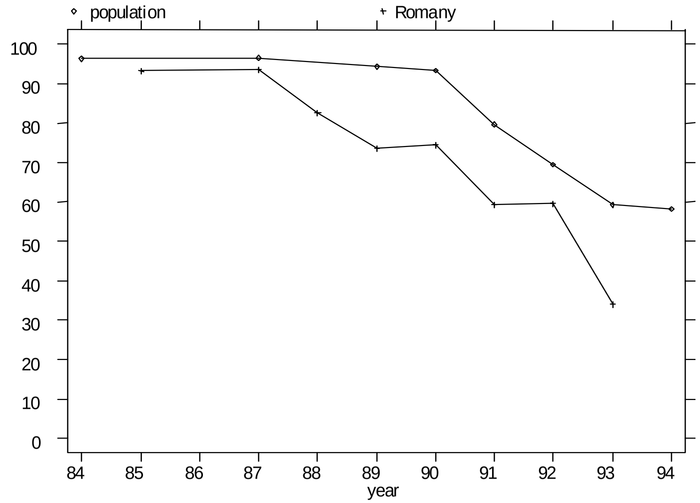
(c) Romany with completed vocational training school

(d) population with completed vocational training school

Graph 9: What percentage of the given age group kept their jobs or found a job between 1984 and 1994 depending on schooling, ethnic origin and gender?

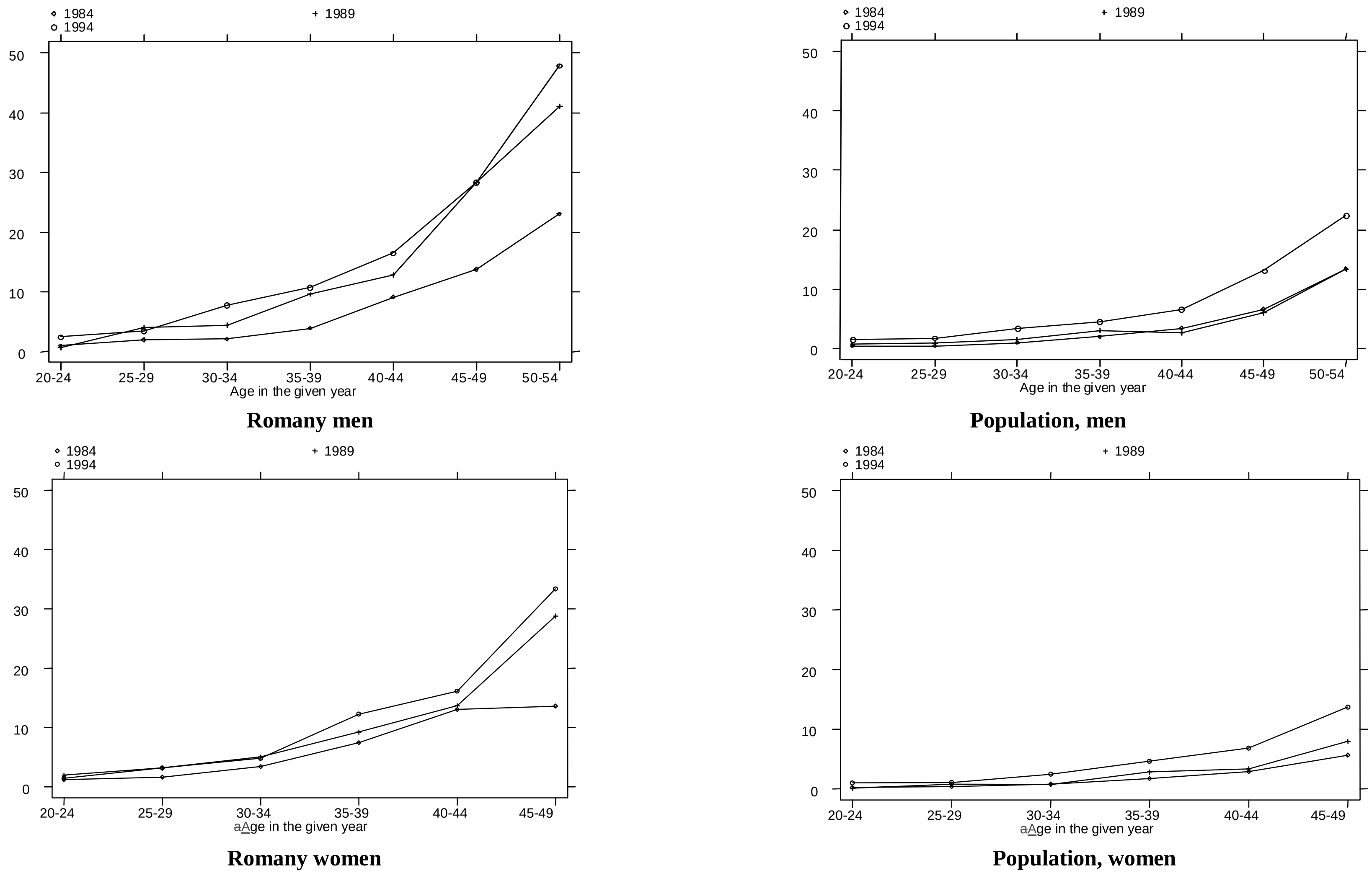


(a) persons with 8 years of education

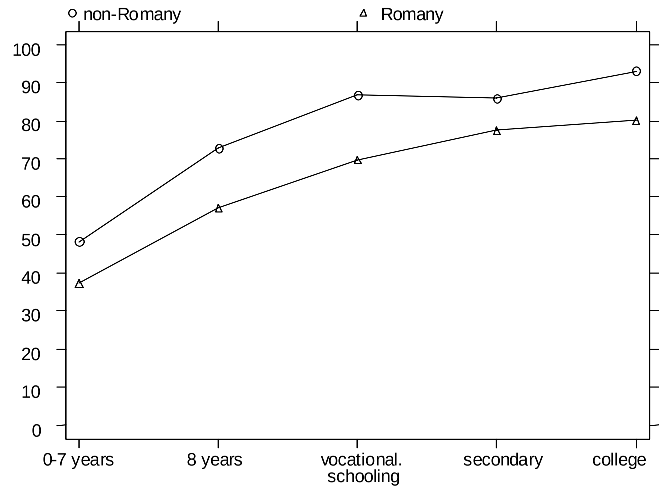


(b) persons with completed vocational training school

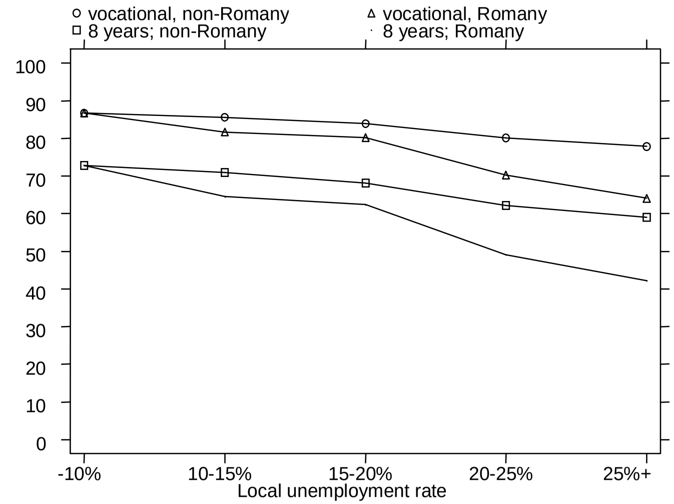
Graph 10: What percentage labour market entrants (aged 15-19) with 8 years of education or completed vocational school could find a job between 1984 and 1994 depending on ethnic origin?



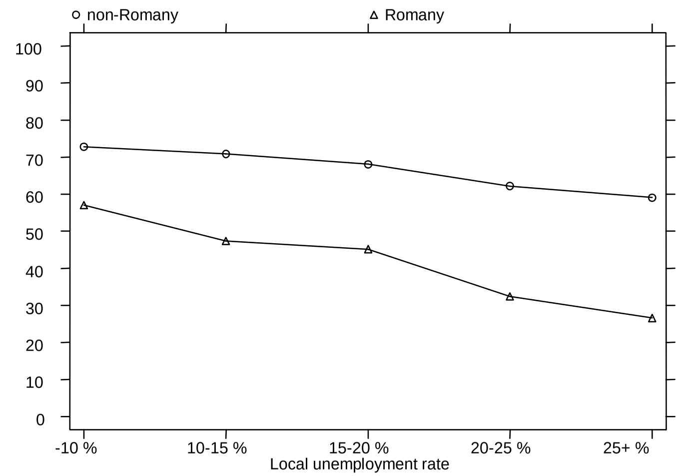
Graph 11: The percentage of retired persons in the given age groups in 1984, 1989, and 1994.



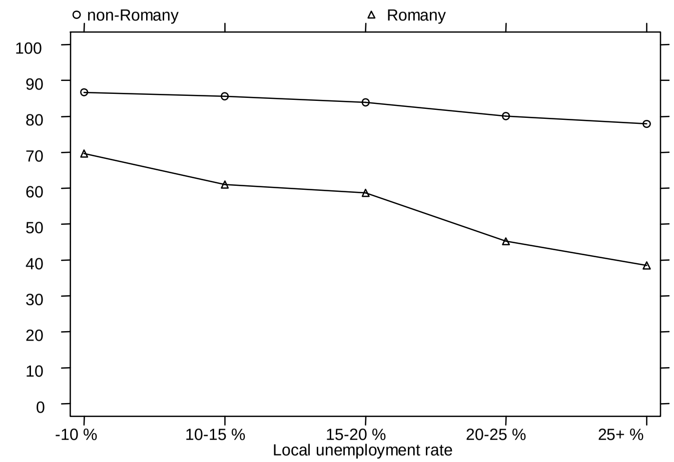
(a) the effect of the constant term and schooling (3)



(b) the effect of the local unemployment rate, with the constant term and the schooling parameters from the non-Romany equation (4*ab*)



(c) the effect of the local unemployment rate, the constant term and schooling parameters for persons with 8 years of education (5*a*)



(d) the effect of the local unemployment rate, the constant term and schooling parameters for persons with vocational education (5*b*)

Graph 12: The predicted employment probabilities of men (based on Tables 10 and 11)

Table 1: Labour market stocks and flows

Year(t)		year (t +1)	
	Employed	Non-employed	Retired
Employed (E_t)	ee_t	en_t	ep_t
Non-employed (N_t)	ne_t	nn_t	np_t
New entrant (Y_t)	ye_t	yn_t	–
All	E_{t+1}	N_{t+1}	P_{t+1}

Table 2: Prior labour market attachment of those Roma workers in 1989, who lost, as opposed to those who managed to keep their jobs by 1994

Age in 1989 (years)	The average number of years worked before 1989 of those, who			
	lost their jobs by 1994		were employed in 1994	
	men	women	Men	women
20 – 24	5.3	5.3	5.8	5.4
25 – 29	9.9	7.9	10.5	8.6
30 – 34	14.3	11.9	14.2	12.4
35 – 39	18.1	13.6	19.7	15.1
40 – 44	23.4	15.0	24.3	17.2
45 – 49	28.9	18.4	28.2	23.4

Table 3: The educational attainment of working age Romany and non-Romany population by gender, 1993 (%)

Education	Men			Women		
	Non-Romany	Romany	Difference	Non-Romany	Romany	Difference
0-7 classes	3,08	30,92	–27,84	2,24	43,46	–41,22
8 classes	31,19	50,45	–19,26	37,79	48,16	–10,37
Vocational school	32,36	16,44	15,92	17,47	6,63	10,84
Secondary school	23,22	1,92	21,30	32,07	1,53	30,54
College	10,15	0,26	9,89	10,44	0,22	10,22
All	100,00	100,00	–	100,00	100,00	–

Note: working age= men: 15-59 years of age, women: 15-54 years of age; population not in school.

Table 4: The employment-population ratio in the Romany and non-Romany population with completed primary and vocational training school, by gender and age, 1993 (%)

Group	age: 15-19	age: 20-24	age: 25-29	age: 30-39	age: 40-54
			men with completed primary school (8 classes)		
Non-Romany	41,7	60,7	66,9	68,8	63,8
Romany	18,2	36,6	38,3	35,1	33,6
Difference	23,5	24,1	28,6	33,7	30,2
			women with completed primary school (8 classes)		
Non-Romany	35,1	30,7	41,0	59,7	58,2
Romany	12,0	11,4	16,3	26,0	30,8
Difference	23,1	19,3	24,7	33,7	27,4
			men, vocational training school		
Non-Romany	53,2	73,0	83,9	79,5	74,9
Romany	23,9	41,2	52,6	50,0	50,8
Difference	29,3	31,8	31,3	29,5	24,1
			women, vocational training school		
Non-Romany	71,6	49,6	44,0	67,8	75,2
Romany	38,8	31,4	33,3	36,9	.
Difference	32,8	18,2	10,7	30,9	.

Note: persons not in school.

Table 5: The distribution of working-age Romany and non-Romany population by type of settlement, 1993 (%)

Group	Budapest	county capital	other town	village	All
Non-Romany	20,21	17,56	26,21	36,02	100,00
Romany	8,02	9,86	19,04	63,07	100,00
Difference	12,19	7,70	7,17	–27,05	–

Note: working age= men: 15-59 years of age, women: 15-54 years of age; population not in school.

Table 6: The distribution of working-age Romany and non-Romany population by the local unemployment rate, 1993 (%)

Group	Local unemployment rate					
	– 10 %	10-15 %	15-20 %	20-25 %	25 % +	All
Non-Romany	32,43	39,93	19,14	6,79	1,71	100,00
Romany	16,37	27,56	32,98	13,20	9,89	100,00
Difference	16,06	12,37	–13,84	–6,41	–8,18	–

Note: working age= men: 15-59 years of age, women: 15-54 years of age; population not in school; local unemployment rate: the unemployment rate of the labour office district, 1993 Autumn.

Table 7: The distribution of working-age Romany and non-Romany population by the local unemployment rate and settlement type, 1993 (%)

Group	Local unemployment rate					
	– 10 %	10-15 %	15-20 %	20-25 %	25 % +	All
			county capitals			
Non-Romany	17,62	73,27	9,11	.	.	100,00
Romany	16,26	68,29	15,45	.	.	100,00
Difference	1,36	4,98	–6,34	.	.	–
			other towns			
Non-Romany	14,42	42,75	31,01	10,90	0,92	100,00
Romany	8,75	23,67	36,73	20,78	10,07	100,00
Difference	5,67	19,08	–5,72	–9,88	–9,15	–
			villages			
Non-Romany	14,84	44,01	26,14	10,92	4,08	100,00
Romany	8,05	25,87	38,78	14,66	12,63	100,00
Difference	6,79	18,14	–12,64	–3,74	–8,55	–

Note: working age= men: 15-59 years of age, women: 15-54 years of age; population not in school; local unemployment rate: the unemployment rate of the labour office district, 1993 Autumn.

Table 8: The employment-population ratio in the working-age Romany and non-Romany population by the local unemployment rate and settlement type, 1993 (%)

Group	Local unemployment rate				
	– 10 %	10-15 %	15-20 %	20-25 %	25 % +
			Budapest		
Non-Romany	64,3	.	.	.	.
Romany	35,8	.	.	.	.
Difference	28,5	.	.	.	.
			county capitals		
Non-Romany	66,5	63,5	59,0	.	.
Romany	31,8	24,9	12,7	.	.
Difference	34,7	38,6	46,3	.	.
			other towns		
Non-Romany	69,5	62,3	60,9	54,9	55,8
Romany	30,2	26,0	23,8	12,4	21,0
Difference	39,3	36,3	37,1	42,5	34,8
			villages		
Non-Romany	65,9	57,4	55,2	47,5	48,7
Romany	36,2	25,0	24,5	16,7	10,9
Difference	29,7	32,4	30,7	30,8	37,8

Note: working age= men: 15-59 years of age, women: 15-54 years of age; population not in school; local unemployment rate: the unemployment rate of the labour office district, 1993 Autumn.

Table 9: The ratio of over-aged and year-repeating students among Romany and non-Romany children attending primary school in 1974/75, 1981/82, and 1985/86.

Group	Class	School year 1974/75	1981/82	1985/86
			Ratio of over-aged students (%)	
Romany students	1-4. class	55.6	41.6	46.7
Non-Romany students	1-4. class	7.4	6.0	9.2
Romany students	5-8. class	62.9	52.7	51.2
Non-Romany students	5-8. class	12.5	8.3	9.3
			Ratio of year-repeating students (%)	
Romany students	1-4. class	22.3	16.3	17.4
Non-Romany students	1-4. class	1.8	1.7	2.4
Romany students	5-8. class	14.5	13.5	14.4
Non-Romany students	5-8. class	1.6	1.5	1.9

* *Source:* Cigány tanulók [1986], pp. 51 and 58.

Table 10: The estimation of employment probabilities (logit)
(men aged 15-59, women aged 15-54, persons not in school)

	Non-Romany*			Romany**		
Independent variable	Coefficient	t -value	p-value	Coefficient	t -value	p-value
Male	–0,409	–14,83	0,000	–0,635	–8,32	0,000
Years of age:						
15-19	–2,315	–40,09	0,000	–0,716	–5,24	0,000
20-24	–1,077	–22,66	0,000	–0,341	–2,98	0,003
25-29	–0,684	–14,40	0,000	–0,157	–1,35	0,177
40-54	–0,185	–4,95	0,000	–0,140	–1,34	0,181
55-59	–1,654	–23,18	0,000	–1,170	–3,71	0,000
Schooling:						
0-7 classes	–1,056	–11,98	0,000	–0,801	–8,83	0,000
Vocational school	0,894	25,00	0,000	0,548	5,21	0,000
Secondary school	0,816	24,11	0,000	0,948	3,93	0,000
College	1,606	28,55	0,000	1,103	1,73	0,084
Married	0,240	6,88	0,000	0,142	1,49	0,135
Number of children	–0,259	–15,80	0,000	–0,202	–7,37	0,000
Local unemployment rate						
10-15 %	–0,095	–3,03	0,002	–0,387	–3,64	0,000
15-20 %	–0,227	–5,99	0,000	–0,477	–4,61	0,000
20-25 %	–0,489	–9,07	0,000	–1,021	–7,16	0,000
25+ %	–0,618	–6,29	0,000	–1,299	–7,84	0,000
Constant term	0,984	18,39	0,000	0,283	1,96	0,050
Log-likelihood	–17483,9			–2265, 4		
LR chi2 (15)	8060,52			528,62		
Pseudo R^2	0,1873			0,1045		
Number of cases	32235			4607		

* CSO Labour Force Survey, 1993 Autumn.

** The 1993/94 representative Roma Survey of the Institute of Sociology, Hungarian Academy of Sciences.

Table 11: The difference in predicted employment rates[a] of Romany and non-Romany men of age 30-39 and the decomposition of of these differences based on alternative assumptions about the source of the employment probability differentials

	the source of the employment probability differentials[b]			The overall difference in predicted employment rates (%)	The overall difference in predicted employment rates= 100%			
					1. decomposition[c]		2. decomposition[d]	
	Constant term	Education	Local unemployment rate		Composition effect (%)	Parameter effect (%)	Composition effect (%)	Parameter effect (%)
1.	(n, r)	(n, n)	ref.	26,3	47	53	61	39
2.	(n, n)	(n, r)	ref.	12,2	100	0	85	15
3.	(n, r)	(n, r)	ref.	26,9	46	54	51	49
4a.	(n, n)	(n, n | e = 8 class.)	(n, r)	8,9	20	80	37	63
4b.	(n, n)	(n, n | e = vocat.)	(n, r)	6,3	20	80	44	56
5a.	(n, r)	(n, r | i = 8 class.)	(n, r)	25,5	7	93	13	87
5b.	(n, r)	(n, r | i = vocat.)	(n, r)	27,6	5	95	14	86

[a] The predictions were based on parameters in the employment probability equations ran separately for the Romany and non-Romany population (see *Table 10*). For the detailed description of the decomposition procedure, see the *Appendix*. The meaning of the (n, r); (n, n), and ref. symbols is the following: (n, r) = for calculating the non-Romany and Romany employment probabilities we used the parameters of the equation of the given population; (n, n) = we used the parameters from the non-Romany equations to predict both the Romany and the non-Romany employment probabilities; ref.= we fixed the value of the given variable at the reference level.

[b] The numbers used in the lines of the table are the same as the decomposition equations in the Appendix.

[c] The composition effect was calculated using the non–Romany parameters.

[d] The composition effect was calculated using the Romany parameters.

Earnings, Age, Education and Ethnicity in the Baltic States

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Abstract

Evidence indicates that labor earnings distributions have undergone substantial shifts vis-à-vis ethnic groups (native Baltic ethnicity and ethnic Russians in particular) in Estonia and Latvia during the transition process. Interestingly, this shift appears not to have occurred in Lithuania, where ethnic conflicts have been largely absent. Since schooling is often segregated by language, and hence ethnicity, in the Baltic States, empirical evidence is examined to determine the extent to which returns to human capital - particularly schooling - vary across ethnic groups. Using data from labor force surveys in the three Baltic States, evidence indicates native Balts have far larger financial returns to schooling in Estonia and Lithuania (despite a relatively small overall earnings gap in Lithuania) while differences in returns to schooling are relatively small in Latvia despite a large overall ethnicity earnings gap. As with education, results indicate no difference in returns to age (experience) in Latvia or to experience and tenure in Lithuania. However evidence indicates higher returns to age for ethnic Russians in Estonia.

Introduction

Several studies have found significant labor earnings differentials across ethnic groups – particularly individuals of native Baltic ethnicity and ethnic Russians – in Estonia and Latvia (see Kroncke and Smith (1999) and Noorkoiv et al. (1998) as examples regarding Estonia and Chase (2000) as an example regarding Latvia). These studies indicate that, controlling for various factors, ethnic Estonians and ethnic Latvians tend to have higher earnings than ethnic Russians in Estonia and Latvia respectively. Evidence on Lithuania remains sparse, though existing evidence provides little indication of significant earnings differentials (Smith (2003)). Further, what evidence exists from the late Soviet period (Smith (2003)), indicates a substantial shift in relative earnings across ethnic groups since the beginning of transition in all three Baltic States.[1]

Despite the evidence on relative earnings across ethnic groups, little evidence exists regarding differential returns to human capital across ethnic groups in the Baltic States.[2] Since human capital is postulated to play such an important role in determining labor earnings, this represents a major shortcoming with respect to understanding earnings differentials across Baltic ethnic groups. Preliminary evidence presented below examines how returns to two key human capital components, education and experience (proxied by age in the Estonian and Latvian Labor Force Survey data), differ between ethnic Russians and ethnic Balts.

Summary of Data, Methodology and Results

Data used in the following estimations are from Baltic Labor Force Surveys (LFS) conducted independently in each of the three Baltic States by their respective Statistical Offices. The estimations for each country use a single cross-section of data. The results thus examine factors affecting earnings at a single point in time. The data for Estonia, Latvia, and Lithuania were respectively collected in January 1997, May 1998, and May 1999.

For all three countries ordinary least squares (OLS) regression is used to examine differential returns to human capital factors and other potential determinants of labor earnings. The results should be viewed as preliminary and suggestive. At this stage, little has been done to measure the sensitivity of the results to alternative specifications of the wage equation and little has been done to test the robustness of the results. Further, given earnings are measured by discrete category rather than as a continuous variable in the Latvian and Lithuanian LFS data, there are potential problems associated with the use of OLS. These problems will be addressed in future work with the data.

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[1] Though in Lithuania the relative shift does not appear to have occurred vis-à-vis ethnic Lithuanians and Russians, but rather through other minority groups such as Ukrainians and Poles.

[2] In the economics literature, human capital typically refers to attributes that will enhance one's productivity as a worker and consequently one's earning ability. Key elements of human capital might include education, training, overall work experience, and tenure on a particular job. The data used here allow for measurement of education, age (to be used as a proxy for experience), or work experience and job tenure (in the case of Lithuania).

Two OLS equations are estimated for each country. The standard method of interpreting the results for education and experience (age) in the presence of interaction terms (see Table 3 for a definition of variables including the interaction terms) are presented below. The β terms represent the OLS coefficient estimates for specific indicated variables presented in Table 2.

First Column for Each Country in Table 2

- Education: $\beta_{education} + \beta_{education*ethnicity}$*ethnicity
- Ethnicity: $\beta_{ethnicity} + \beta_{education*ethnicity}$*education
- Age: $\beta_{age} + 2\beta_{age2}$*age

Second Column for Each Country in 2

- Education: $\beta_{education}$
- Ethnicity: $\beta_{ethnicity} + \beta_{age*ethnicity}$*age
- Age: $\beta_{age} + 2\beta_{age2}$*age + $\beta_{age*ethnicity}$*ethnicity

A nontechnical discussion of results for each country follows.

Estonia

Table 1 presents descriptive statistics for the three countries. The Estonian income figures indicate a substantial earnings gap between ethnic Estonians and ethnic Russians with ethnic Russians earning about 84 percent of what Estonians earn on average. As Table 1 further indicates, average educational level and average age are quite similar for the two ethnic groups. Though not presented, the distribution of ages and educational levels are also quite similar for the two groups (This holds true for Latvia and Lithuania as well).

A simple OLS regression controlling for gender, age, education and location in major urban centers (results are not presented though available upon request from the author as are results for regressions on Latvia and Lithuania that exclude interaction terms) indicates that nearly 90 percent of the earnings gap between Estonians and Russians can be attributed to ethnic background. Table 2 presents OLS results indicating relatively standard gender earnings differentials. The results predict, given existing controls, that women can expect to earn about 28 percent less than men. Further the results indicate relatively significant and strong returns to education and a somewhat normal (relative to most market economies) age-earnings profile for workers. The age results do indicate a fairly early peak in earnings – at about 42 years of age – as opposed to most market economies (where earnings are more likely to peak in the early to mid-50s) though one that seems fairly typical of transition economies.

Table 2 also presents results that allow for a specific examination of differential returns to education and age between the two ethnic groups represented in the sample. The first column of Table 2 focuses on differential returns to education between ethnic Estonians and ethnic Russians. The results indicate a statistically and practically large ethnic effect. The predicted return to attaining a higher level of education (the Estonian data separates individuals into 7 educational categories) is roughly 14.2 percent for Estonians (found by summing the coefficient estimate for education and the coefficient estimate for the interaction term, education*ethnicity) as opposed to roughly 6.5 percent for Russians.

The results also indicate a relatively small earnings gap favoring ethnic Estonians for those who have the lowest level (primary only) of educational attainment. Interpreted literally, the results predict Estonians who have not advanced beyond a primary education will earn five percent more than Russians who have not advanced beyond a primary education. However, the gap favoring ethnic Estonians widens quickly at higher levels of education. The results further predict that an Estonian with a bachelor's degree, all else equal, would earn approximately 35 percent more than an ethnic Russian with a bachelor's degree.

With respect to age the results presented in column 2 of Table 2, indicate a wide earnings gap favoring ethnic Estonians when young that shrinks as workers age. The regression predicts a 20-year old Estonian worker can expect to earn about 28 percent more than a 20-year old Russian worker controlling for other factors. The predicted gap between the groups gradually shrinks for older workers. For two workers aged 60, the results predict the Estonian worker will earn only about eight percent more than the Russian worker. As a note, given the cross-sectional nature of the data, these results do not indicate that the ethnic earnings gap that exists between young workers will shrink over time. It may well be that the large ethnic earnings gap existing between contemporary young workers in Estonia will persist throughout their working lives implying a very large lifetime earnings gap.

Latvia

While evidence exists indicating significant earnings gaps favoring ethnic Latvians vis-à-vis ethnic Russians, no evidence is presented here indicating a human capital explanation. A simple regression (not presented) that excludes interaction terms but includes the other controls indicated in Table 2, does predict that Latvians on average will occupy a higher position in the overall Latvian earnings distribution than will ethnic Russians. The results also indicate that gender, age and education affect earnings in fairly standard ways by the norms of a market economy (as in Estonia with a fairly early peak in the age-earnings profile). However, OLS results indicate similar returns to education and similar age-earnings profiles between the two groups. Clearly one must look elsewhere to explain the ethnic earnings gap that appears to exist in the Latvian labor market.

Lithuania

The general situation regarding the ethnic Russian minority is quite different in Lithuania as opposed to Estonia and Latvia. As Table 1 indicates, the ethnic Russian minority in Lithuania is quite small and represents only about seven percent of the Lithuanian LFS sample used in the estimations. Further, a basic OLS regression excluding interaction terms does not provide any evidence of earnings differentials between ethnic Lithuanians and ethnic Russians. The Lithuanian regressions do however provide evidence of a large gender earnings gap and very strong returns to education. Due to the structure of the Lithuanian LFS, overall work experience and job tenure are used in the OLS regressions rather than age. The results (quite similar to those in Table 2) with respect to experience and tenure are quite interesting. The results for experience imply that general work experience does not significantly affect earnings. However, the tenure result indicates experience on the current job has a statistically significant and practically important influence on earnings. Specifically the numbers imply that tenure initially increases earnings though the effect fades as tenure lengthens. Given the transitional state of the Lithuanian economy, this is not entirely surprising. A reasonable explanation might be that job specific skills – particularly those gained in the post-Soviet period – significantly increase productivity and consequently labor earnings while general work experience – perhaps largely gained in Soviet era enterprises – is not perceived as valuable in a market economy.

Turning to Table 2, with respect to education, Lithuania is somewhat similar to Estonia. The results indicate much stronger returns to higher educational attainment for ethnic Lithuanians as opposed to ethnic Russians. Additionally the results provide evidence that ethnic Russians with low levels of education fare considerably better (given the Lithuanian survey, this implies a better standing on average in the overall earnings distribution) than ethnic Lithuanians with equally low levels of education. However Lithuanians catch up quickly as educational level rises and at the highest levels of educational attainment (those with higher education), the evidence indicates ethnic Lithuanians on average have a higher standing in the earnings distribution than do ethnic Russians. As opposed to Estonia though, the results of the final column of Table 2 indicate that experience has little influence on the relative earnings distribution vis-à-vis ethnic groups. Though not shown in Table 2, the same appears to be true of tenure.

Concluding Remarks

This study represents a preliminary attempt to examine how certain human capital attributes affect earnings in the three Baltic States and the extent to which returns to these human capital factors are different with respect to those of native Baltic ethnicity and Russian ethnicity. Baltic LFS data provide evidence of substantial earnings gaps between ethnic Balts and ethnic Russians in both Estonia and Latvia. However no such evidence of an ethnic earnings gap exists in Lithuania.

The regressions discussed above provide evidence of differential returns to education in Estonia and Lithuania and an ethnic earnings gap that declines with worker age in Estonia. Despite evidence of a substantial earnings gap in Latvia, the human capital factors examined here do not seem to contribute to the gap.

To the extent the results have implications for policy, the implications would tend to be most clear (though hardly definitive at this point) for Estonia. Given that a significant portion of the overall Estonian ethnic earnings gap appears to be attributable to differential returns to education, and given that education in the Baltic States is frequently segregated by language, it would seem efficacious to either concentrate resources on improving the quality of Russian language education or improving the access of ethnic Russians to broader educational opportunities.

Table 1
Descriptive Statistics
(Labor Force Survey Data – 1997-1999)

	Estonia			Latvia			Lithuania		
variable	pooled sample	Estonians	Russians	pooled sample	Latvians	Russians	pooled sample	Lithuanians	Russians
income	2663.7 (2029.8)	2790.2 (2188.1)	2338.6 (1503.7)	2.950 (0.997)	2.947 (1.001)	2.956 (0.986)	3.643 (1.996)	3.605 (2.007)	4.154 (1.768)
education	2.532 (1.524)	2.522 (1.540)	2.554 (1.481)	5.985 (1.607)	5.989 (1.633)	5.974 (1.543)	5.151 (2.026)	5.129 (2.037)	5.435 (1.854)
age	40.985 (12.142)	41.133 (12.297)	40.604 (11.735)	39.641 (11.795)	39.641 (12.114)	39.639 (11.005)	--	--	--
experience	--	--	--	--	--	--	18.585 (11.590)	18.525 (11.699)	19.378 (10.021)
tenure	--	--	--	--	--	--	7.248 (8.274)	7.244 (8.271)	7.298 (8.336)
N	2492	1794	698	4887	3439	1448	3496	3250	246

Notes:

a) Mean values are given with standard deviations in parentheses. N represents the number of observations in each sample. The Baltic LFS data includes other ethnic groups as well. Samples are restricted to include only native ethnic Balts and ethnic Russians.
b) All variables are defined in Table 3.

Table 2
OLS Regression Results
(Labor Force Survey Data – 1997-1999)

variable	Estonia		Latvia		Lithuania	
intercept	6.505*** (0.146)	6.237*** (0.155)	1.833*** (0.164)	1.784*** (0.155)	2.257*** (0.210)	1.526*** (0.139)
ethnicity	-0.027 (0.059)	0.382*** (0.105)	-0.024 (0.113)	0.031 (0.105)	-0.812*** (0.204)	0.024 (0.143)
gender	-0.280*** (0.026)	-0.279*** (0.026)	-0.413*** (0.027)	-0.413*** (0.027)	-0.716*** (0.051)	-0.715*** (0.051)
age	0.050*** (0.007)	0.052*** (0.007)	0.029*** (0.007)	0.029*** (0.007)	--	--
age^2	-0.0006*** (0.00008)	-0.0006*** (0.00008)	-0.0003*** (0.00009)	-0.0003*** (0.00008)	--	--
experience	--	--	--	--	-0.006 (0.008)	-0.005 (0.010)
$experience^2$	--	--	--	--	0.0004** (0.0002)	0.0004** (0.0002)
tenure	--	--	--	--	0.048*** (0.009)	0.046*** (0.010)
$tenure^2$	--	--	--	--	-0.0006** (0.0003)	-0.0006* (0.0003)
education	0.065*** (0.0168)	0.121*** (0.009)	0.158*** (0.016)	0.174*** (0.008)	0.315*** (0.034)	0.457*** (0.014)
education* ethnicity	0.077*** (0.019)	--	0.021 (0.018)	--	0.166*** (0.037)	--
age*ethnicity	--	-0.005** (0.002)	--	0.002 (0.003)	--	--
exp*ethnicity	--	--	--	--	--	0.001 (0.006)
urban controls	yes	yes	yes	yes	yes	yes
F statistic	53.64***	52.30***	134.43***	134.29***	184.34***	181.81***
R^2	0.178	0.174	0.162	0.162	0.383	0.380
N	2492	2492	4887	4887	3496	3496

Notes:

a) The dependent variable is log(earnings) for Estonia and earnings for Latvia and Lithuania. Individuals place themselves in earnings categories in the Lithuanian and Latvian surveys.

b) Given the earnings definition in the Estonian data, interpreting the Estonian results is fairly straightforward. As an approximation, the coefficient estimates can be multiplied by 100 to get a percentage effect on earnings. For example, given the definition of gender, a coefficient of –0.28 indicates, controlling for other variables in the regression, women in Estonia are predicted to earn 28 percent less than men. Unfortunately, given the earnings definitions in the Latvian and Lithuanian samples, interpretation of results is less straightforward and there are problems associated with OLS estimation. However, the problems are unlikely to affect the qualitative interpretation of the results.

c) Standard errors are in parentheses. *** denotes statistical significance at the 1% level, ** denotes significance at the 5% level, and * denotes significance at the 10% level.

Table 3
Variable Definitions

variable	variable definition
earnings	monthly earnings in EEK in Estonia and by category in Latvia and Lithuania
ethnicity	= 1 if a person is of native Baltic ethnicity and = 0 if ethnic Russian
gender	= 1 if female and = 0 if male.
age	age in years (used for Estonia and Latvia)
age^2	age-squared (used for Estonia and Latvia)
experience	total years of work experience (used for Lithuania)
$experience^2$	experience-squared (used for Lithuania)
tenure	total years of experience on the current job (used for Lithuania)
$tenure^2$	tenure-squared (used for Lithuania)
education	Ascending levels of educational attainment
education* ethnicity	an interaction term multiplying ethnicity and education
age*ethnicity	an interaction term multiplying ethnicity and age (experience for the Lithuanian regressions in Table 2)
urban controls	dummies for certain urban areas (i.e., capital cities)

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The Impact of Brain Drain on East German Workers*

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1 Introduction

After unification of Germany in 1989, a huge discussion about mobility in Germany started. Cumulative net migration flows from East into West Germany amount to almost 1.3 million people during the time from 1989 until the end of 2001. This corresponds to a share of 7.5 per cent of the 1989 population in East Germany. Although net migration rates in the second half of the 1990s are much below those of the initial years, they have accelerated again after 1996. This increase coincides with the end of the convergence of per capita income levels between the West and the East of Germany. Also income inequality stopped to converge in 1996 (see figure 1).

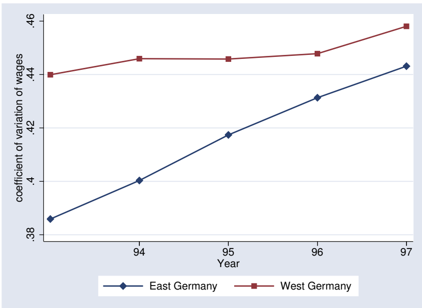

Figure 1: Inequality in East and West Germany

The persistent phenomenon of East-West migration in Germany has raised increasing concerns that workers with the highest qualifications and the highest abilities move to the West and that this "brain drain" will further contribute to sluggish economic growth in East Germany and divergence of per capita income levels between East and West Germany.

But not only the economic discrepancy between East and West Germany is striking, also the regional discrepancies in East Germany are considerable. This economic inequality leads to a highly unequal distribution of labour market perspectives. Several reports show the unequal distribution in East German Länder or districts and also try to find ways to help the most backward regions (see for example Blien, Blume, Eickelpasch, Geppert, Maierhofer, Vollkommer, and Wolf (2001). As the development of East German districts is diverse and also the emigration from the districts is not uniform over East Germany, we want to look deeper into the structure of movements especially of high skilled workers and the resulting labour market developments in East German districts.

The reminder of the paper is as follows: Chapter two contains a short overview over existing brain drain theories to help finding hypotheses. The different data sets used are described in chapter tree, chapter four shows first descriptive evidence for brain drain in East Germany. In chapter five we attempt to explain the economic performance of stayers in East Germany with the emigration of high skilled labour.

2 Theoretical foundation

The development of brain drain theories had its first peak in the 60'ies when migration of skilled people from developing to developed countries accelerated, but there are still developments of theories going on today that help to explain and understand brain drain phenomena. Brain drain means the migration of skilled labour from one region or country to another one because of economic differences between these two regions. The discussion has always been whether this brain drain leads to welfare losses or gains in the sending region. Theories in this chapter are described following the paper by Commander, Kangasniemi, and Winters (2003).

2.1 Brain drain theory

2.1.1 Early work

Early brain drain theory models the labour market of the sending country using static analyses. Grubel and Scott (1966) develop one of the earliest models of brain drain supposing a perfectly competitive market. This model leads to no welfare impact of skilled emigration on those left behind, because wages are set to marginal productivity and there do not exist any externalities. Also markets always clear.

Later papers introduce several distortions to make a welfare loss for the economy possible. The most important distortions introduced by subsequent papers where that there could exist a gap between social and private marginal product and there is publicly subsidised education. The second is important when highly skilled people leave the country without giving back the education they received in the form of high productivity.

Bhagwati and Hamada (1974); Hamada and Bhagwati (1975) use a general equilibrium model to model the influence of high skilled emigration on stayers and on the sending country. Two kinds of distortions are introduced: A special wage setting procedure and the financing of education. High skilled wages are determined via international emulation which means that they are determined partly by foreign skilled wages. This is indeed the case when two economies integrate, because skilled labour can easily move to the region with the higher wage. Unskilled wages are determined by "leap-frogging" meaning that they rise with rising wages for highly skilled.

The result of this model is that skilled emigration can influence the wages in both sectors and also expected wages and education decisions. In a whole, the model predicts a welfare loss of the sending region, the distortions of the labour market are worsened by the loss of skilled workers. Unemployment may raise because of raising wages and education costs may raise because of higher expected wages in the receiving region. Looking at the whole population including the emigrants a welfare gain cannot be excluded because of the gain of the emigrants. The sending country itself is more likely to have a welfare loss.

Further channels to generate a positive effect of brain drain are remittances to the stayers, return migration of individuals with new skills acquired abroad and the creation of business networks between stayers and movers.

2.1.2 Later work

Later models use dynamic specifications and model the negative effects of brain drain for the sending country in an endogenous growth framework (for example Wong and Yip (1999)). Others focus on the motivation of emigration possibilities for human capital accumulation (Mountford, 1997; Vidal, 1998; Beine, Docquier, and Rapoport, 2001). These are the most optimistic models regarding the effects of brain-drain on the sending region.

The motivation to move is again not endogenous in the model. An individual will move if it can profit from the move and it will stay if moving generates a loss. This means, skills have to be rewarded higher in the receiving region to motivate individuals to move at all.

The effect of migration is not clear from the model. Emigration of high-skilled per se is negative, but there can exist encouragement effects on those left behind to accumulate more skills when free movement is installed. This could even lead to an overcompensation of the negative effect of brain drain. If in top some mechanism occurs that generates also gains to others than the skill-accumulating person, an even higher positive effect of brain-drain can occur. These mechanisms comprise for example spill-overs between skilled workers (postulating that the productivity of the workforce depends on the educational achievement one period before) (Mountford, 1997) or intergenerational transmission of skills and education (Vidal, 1998), postulating that the next generation can create skills more easily the higher the skill creation of the leading generation has been.

A strong assumption of the model to generate positive effects of skill accumulation in the sending region is that there must be high-skilled that do not move when they have finished to accumulate skills. This means that there must exist a mechanism to keep some high-skilled in the sending region. This can be modelled by introducing an exogenous probability to move that is smaller than one when an individual is high-skilled. Or the model can introduce the mechanism that every individual moves with a positive probability because firms in the receiving region are not able to screen workers perfectly. Then, some skilled individuals will have to remain in the sending region while some less skilled will move.

With perfect screening of the skills of possible migrants, the receiving country will only employ the most able competitors. In this case, that marginal student will not change his education decision and the positive effect of skill creation does not happen.

Another positive effect can arise when market failures in the sending region exist. When there exists ex-ante unemployment among the high-skilled, welfare gains can be generated. The abilities of the left behind can be used more efficiently when some of the high-skilled move to the receiving region.

2.2 New Economic Geography

New Economic Geography models introduced by Krugman (1991) are general equilibrium models. They model the dynamics of industrial concentration when economies of scale and transportation costs are present. Even when two countries start with the same endowment with industries, economies of scale can lead to uneven development of the two regions. An immobile sector makes complete specialisation impossible, economies of scale drive firms of one sector to move in one of the regions. Labour follows and increases demand for goods, that is served by firms nearby because of transportation costs.

The motivation of high-skilled to move from one region to the other arises if the sector that agglomerates in one region is the skill-intensive high-tech sector. By concentrating in one region, the industry also needs the high-skilled to concentrate in this region and brain drain from the sending (remote) region is the consequence.

From new economic geography models, we can conclude that unequal development in two regions is a natural outcome during agglomeration and that this agglomeration becomes more likely, when trade costs fall – and with them transport costs. The second important conclusion to mention is the most important for our investigation. During agglomeration, differing real wages of regions are normal, because the winning region can generate a higher productivity of (high-skilled) labour. Thus, skilled labour moves not only because the industry is allocated mostly in one region, but also because it pays off to move in the area with the higher real wages.

This means that brain drain leads to a negative effect on welfare of the sending region even in the absence of labour market failures that were the leading factor for negative effects on the sending regions in the former models. Also, there is no mechanism to generate a reverse effect like return migration etc. The only possibility for positive effects on the sending region could be falling transaction costs so that the left-behinds would profit from lower prices of goods produced in the receiving (agglomeration) area.

3 Used Datasets

3.1 Individual Data

We perform our empirical analysis using individual data from the "IAB-Regionalstichprobe".[1] This data set contains a five per cent sample of all the returns of the social security files of East Germany, collected by the Federal Employment Services (Bundesanstalt für Arbeit). The East German sample starts at the beginning of 1992 and the last spells are reported for 1997.

The sample covers employed persons, unemployed persons and individuals who are currently taking a break from employment. Self-employed persons and those who are enrolled in educational programs are not included. Moreover, the sample is censored from above, i.e. individuals whose earnings exceed the rather high ceiling for contributions to the public pension scheme and unemployment insurance in Germany are not reported.[2] In 1995, 86.2% of the economically active population was captured by the social security files in East Germany (Bender, Haas, and Klose, 2000, p. 3).

The observations of each individual are organised as event data. Every change in the employment situation is collected with the date of its event, but also every year a control return is registered. For each individual, work history, personal characteristics, firm characteristics and regional details are collected. We choose only individuals who are employed full-time on 31 March. The employment state on 31 March of every year is used to transform the event-oriented data into a panel of yearly observations.

Only East German observations (workplace, not place of residence) are taken into account in our regression, because we are interested in the impact

[1]Employee sample, regional file. The IAB-Regionalstichprobe is provided by the German Institute for Employment Research (IAB) at the Federal Employment Services (Bundesanstalt für Arbeit). See Haas (2001) for a brief introduction.

[2]The ceiling was 5,300 DM in 1992 and 7,100 DM in 1997, while the mean incomes in our sample amount to 2,695 and 3,097 for the two years.

of emigration on the workers left behind. Observations date from 1993 to 1997, 1992 could not be included because the migration variable had to be created with lagged observations.

3.2 Data on Mikrozensusregion (District)-level

One district level dataset is provided by the IAB and is aggregated from the entire social security file containing all employed in East Germany that have to pay social security. We use the number of employed in three sectors of the economy, in the tradable, the non-tradable and the agriculture and mining sector.

The second dataset contains yearly data from the 5 per cent random sample of East Germany, aggregated by ourselfs, for the years 1993 to 1997. A variable created from this dataset is the migration rate of the three different skill groups. This variable is calculated by first counting all the individuals moving from one district to any other (East or West Germany) in every single year and setting this number in relation to all the workers of the district. This is done with out- and in-migration for each district and the average net migration rate per year is taken as the explanatory variable in our regressions.

An explanation of the economic situation of East Germany can be based on different theories. Based on the brain drain argumentation, emigration of high skilled labour can lead to a worse performance of the economy. As a measure for emigration we use first the above defined mean regional migration rates over the time period 93-97 (Figure 8) and second the change in the share of highly educated workers 93-97 (Figure 9). We can see from Figure 8 that high-skilled workers from the south west part that have a relatively short way to the west labour market have a very high propensity to move and the far north east part is loosing it's high-skilled labour.

The third set of variables contains dummies that divide the districts into three agglomeration degrees provided by the BBR (Bundesamt für Bauwesen und Raumordnung). Finally we have data on place of residence and population density for 1993 to 1997, also provided by the BBR.

3.3 Data on Bundesland-level

Data on Bundesland-level are needed when an aggregation over education group and sector is needed because of the small subgroups when taking district levels. We take the data from the IAB-subsample and aggregate over region, education and sector. The finally used variable in this aggregation level is the unemployment rate on the Bundesland/educational group/sector-level (e.g. unemployment rate of highly educated in the non-tradable sector in Brandenburg). Furthermore we use GDP per capita of Länder from the Gutachten des Sachverständigenrats.

4 Descriptive Evidence of Brain Drain

4.1 Population and Employment in East German districts

To test the brain drain story we are firstly interested in the population- and employment growth on district level whereby the employment growth will be the decisive variable regarding labour market influences. Figure 2 shows the yearly growth rate of the population over the years 1993 to 1997, taken from the data set of the BBR. As can be seen clearly, people move out of the cities to the surroundings. The most depopulating districts are the north-east and also the south-west districts, a fact that cannot be interpreted without knowledge of the underlying economic structure of the region.

As mentioned before, employment growth in the districts is a better indicator regarding labour market issues. Figure 3 shows that employment growth develops less smooth than population growth, but with a similar geographic pattern. The far east part of the districts and also the south west are relatively less growing than the middle part. The surprising fact is that employment decreases in cities similar to the population downturn. One would expect that the urban employment grows because of agglomeration effects in East Germany. To get an impression of the differences between East and West Germany, we show also the employment growth of whole Germany in Figure 4.

For a short overview of the employment growth in the different sectors and over different qualifications, Table 1 shows some summary statistics. It is shown that only in the sector agriculture and mining, the yearly employment growth is positive, in tradable and non-tradable good sectors, employment declined over the observed period 1993 to 1997. All qualification groups loose employed over the observed period. The highest decline is in the low qualification sector while the middle qualified show the most stable number of employed over time.

Table 1: Economic Indicators and their variation over East German Districts

	Mean	Std. Dev.	Min	Max
Employment growth – branches 93-97				
Agriculture, Mining	.0022409	.108463	-.1705196	.4838275
Tradable Goods	-.0256673	.0346121	-.1115837	.0514464
Non-tradable Goods	-.0148621	.0236985	-.0539379	.0661397
Employment growth – qualifications 93-97				
low qual.	-.0616593	.0380836	-.1344953	.06207
high qual.	-.02373	.0370958	-.0883321	.1217903
mid qual.	-.0189756	.0196531	-.0608795	.0445518
Wage growth – qualifications 93-97				
all qualifications	.0492352	.0080236	.0266742	.0802004
high qual.	.066898	.0084446	.0399936	.0983204
mid qual.	.0477791	.007412	.03125	.0759972
low qual.	.0458206	.0125022	.0188884	.0833556
Growth of per capita GDP – Bundesland-level 93-97				
wrgdppc	.0775383	.024625	.0303218	.0992268
Data from IAB, District levels				

4.2 Economic situation of the East German districts

In order to measure the economic achievement of East Germany, wage and unemployment growth can be consulted. Figure 6 shows that again the north- and south-east parts develop less fast than the middle part, now with respect to the yearly wage growth. Berlin is a particularly low growing area, but when looking at Figure 5, one can see the very high wage level at the beginning of our period that can explain the lower growth rates. Looking again at Table 1, one can see the wage growth for different qualifications. The "all qualifications" number now differs from figure 6, because in the table, the whole number of employed is used, while for the figure, we used only a five percentage subsample (see chapter on used datasets). Coming to the numbers, we find that high qualified get the highest wage increase while low qualified get the lowest wage increase, but almost not different from the mid qualified. In the aggregate, an almost 5 percent wage increase per year is found which is quite high, but in line with the catch-up process of East German wages to the West German level (which is not reached yet). As a reference number, yearly GDP per capita growth from 1993 to 1997 is shown in table 1. With a growth rate of almost 8 per cent, it is far above the wage growth observed which could lie in the fact that the means are not weighted with population or workforce, so that there can be a bias in the numbers.

Coming to another measure of economic situation, the unemployment development in the districts. Figure 7 shows that the north and the far south are less hurt by unemployment growth while the middle part has a very high growth in unemployment rates.

4.3 Data basis for our regressions

The descriptives of our regression data set is contained in table 3. We use two time periods for our estimation, one from 1993 to 1995 and one from 1995 to 1997. As East German wages start at a relatively low level, the wage growth is much higher in the first time period. The wage growth shrinks from around 9 percent to 4.6 percent in the second period, while the unemployment rate measured on the Bundesland/education/branch-level stays almost constant at around 14 percent. The mean outmigration rate of high qualified workers is astonishingly negative, meaning that we find at the mean immigration in East German regions of around one percent. This fact is contradictory to macro data that tell that people tend to migrate from East to West Germany. The reason for this number could be that the mean is not weighted. All other indicators are included as control variables and do not show significant differences between the two time spans.

4.4 Attrition in the data set

As we do not use macro (aggregated) data, we have to check whether the wage growth numbers per region are biased because of selection problems. If only the least able stay in a region and all others leave, the wage growth would be underestimated due to a selection bias. People moving in with higher wage growth would also not be considered. To keep the selection effect as small as possible, we divided our sample into two time spans, 1993 to 1995 and 1995 to 1997. The numbers are presented shortly in Table 2.

Table 2: Attrition over time

	1993-1995	**in perc.**	**1995-97**	**in perc.**
Employed stayers	23240	0.812	22039	0.791
Stayers unemployed in 2nd yr.	2190	0.077	3063	0.110
Employed movers	2945	0.103	2311	0.083
Movers unemployed in 2nd yr.	240	0.008	444	0.016
sum first year employed	28615	1	27857	1

In both time spans, around 80 per cent of the low and middle qualified are at the end of the period still in their starting region and are still employed, so that we can calculate their wage growth. The remaining 20 per cent divide into people that dropped out of the sample because they became unemployed and people who moved in another district and have found a new employment. People dropping out because of unemployment are around nine percent of the starting sample, with the most of them still being in their starting region. Only around 1 percent of the sample moved and became unemployed (in both periods). And finally the last 10 percent of the sample are movers with a new job at their destination district.

5 Empirical Test of Brain Drain

The results of the test on a brain drain in East Germany are presented in Table 4. For our wage regressions, we use the following equation which is based on the idea of growth theory:

$$\begin{aligned} wagegrowth_{93-97} &= \beta_1 * individual\ characteristics && (1)\\ &+ \beta_2 * District\ characteristics && (2)\\ &+ \beta_3 * Charact.\ on\ Educ.\ Branch\ Laender\ level && (3) \end{aligned}$$

The idea is that prosperity in the district should suffer from outmigration of highly skilled workers. As we do not have any economic prosperity indicator of the districts, we use as a proxy the yearly wage growth of the relatively immobile workers in micro data form. These are all workers except the highly skilled (workers without qualification and workers with training). The first two columns of Table 4 are regressions for the first time span from 1993 to 1995, the third and fourth are regressions for the second time span from 1995 to 1997. To control for endogeneity of the brain drain indicator (details see below), we always show one OLS and one Instrumental variable regression where the brain drain indicator is instrumented with the share of parents in the district. This instrument is taken, because studies on East Germany found out that parental status is not related to wages (in contradiction to for example studies for the U.S.), a finding that we can reproduce with our dataset. The dependent variable is yearly wage growth of unskilled and skilled with a-level degree 1993 to 1995 and 1995 to 1997.

To control for wage growth that is higher than productivity growth, we introduced as a control variable first the unemployment rate in the district, but this variable showed up to be insignificant. The second unemployment variable we introduce is unemployment aggregated on Bundesland/Education group/Branch-level in the first year (1993 and 1995) (*UE Land/Educ/Branch*). This variable shows a significantly positive influence which is in the first place counterintuitive, but it can be due to endogeneity problems. Another argument for finding a positive influence can be that branches with the highest productivity (and therefore wage) growth grow faster because they rationalise their firms and have to lay off more workers than in other branches. As the other coefficients do not change when omitting the variable, we include it without instrumenting for it, because of the lack of good instruments and because it is no key variable in our reasoning.

The important variables to look at are at the bottom of Table 4. We use the mean yearly net emigration rate of highly educated employed over the years 1992 to 1994 and 1994 to 1996 (*Emig. HQ district*) as indicator for a possible brain drain. We find that outmigration influences the wage growth of less skilled in the district negatively. As a further control variable, we use the initial share of highly educated employed (*Share HQ in district*) to see if it helps the district to grow if there exists a higher stock of highly skilled workers. As expected, the higher the share of initial human capital in the district, the higher is the wage growth of the less mobile, our indicator for the prosperity of the district. Blien, Maierhofer, Vollkommer, and Wolf (2002) get in their study a similar result. They find that the growth of employment is higher in districts where the workers are higher skilled than the average.

To avoid biased results regarding the brain drain indicators, we use instrumental variables estimation in columns 2 and 4. Finding an instrument is particularly hard because of the lack of rich data sets and because of the very close theoretical relation between the wage growth and brain drain indicators. As instruments for the emigration rate, we use the share of families with children as a hindrance for migration as explained above. The coefficients of the instrumented variables increase by a huge amount, but they still keep their sign and significance, so that we can conclude with our regressions that there exists a positive relation between skills in production so that the leaving high-skilled deteriorate the economic situation in the district.

The remaining variables are mainly to capture the heterogeneity between the individuals, but they can also be interpreted as in normal wage regressions. We find a standard relation between age and wage growth with diminishing wage growth when workers get older (this is dependent also on the height of the wages that get higher with higher ages). Unskilled (blue collar) workers get the highest wage increase in the regressions, while white collar workers (skilled and clerks and foremen) get less wage increase. Workers in the agriculture and mining sector get the lowest wage increase, while we have seen in table 1, that the employment growth was highest there. This employment growth seems to be bought with a very moderate wage increase in this sector. The highest wage increase was realised by the workers in the tradable goods sector. Marital status does not influence wage growth, while males got only a higher wage growth in the first period, not any more in the second. The last indicator is the agglomeration level, where workers in agglomerated areas got a higher wage increase than workers in urbanised or rural areas which is an expected result.

To test whether this negative effect of outmigration is because of brain drain or just because of losing labour, we also calculated the effect of low skilled emigration on high skilled wage growth, where we could not find a significant result. This affirms that the result is driven by a brain drain effect in East German districts.

6 Concluding remarks

Different to many studies that investigated the impact of brain drain on the sending economies, we find in our investigation, that signs for a bad influence of brain drain exist. Looking at East German districts, workers can realise a higher wage growth if there is a share of high qualified in the same district and a lower emigration rate of high qualified out of this district.

Shortly after reunification of Germany, a huge discussion started about whether it would be more helpful if encourage workers to leave the East or if it would be more helpful if the state tries to keep all workers there. Looking at our results and assuming that high-skilled workers are more mobile, we would conclude that it is better for the districts in the East to keep the whole workforce, not only the immobile. This argument aggravates if one accounts for all the movers that are not in our sample because they move before they start working. This is again true for many high-skilled workers that already study in the west or move after finishing university.

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Table 3: Summary statistics

Variable	**Mean 1993-95**	**Mean 1995-97**
wagegrowth	0.094	0.046
	(0.161)	(0.121)
age	37.752	38.711
	(10.357)	(10.34)
age_2	1532.469	1605.435
	(791.475)	(806.441)
Trainees and unskilled	0.162	0.161
Skilled workers	0.429	0.413
Clerks and foremen	0.409	0.426
Agriculture, Mining	0.085	0.069
Tradable Goods	0.235	0.226
Nontradable Goods	0.68	0.705
Marital status	0.541	0.571
Male	0.586	0.577
Agglomeration area	0.355	0.364
Urbanised area	0.417	0.41
Rural area	0.227	0.226
Unempl. rate Land/Educ/Branch	0.134	0.141
	(0.035)	(0.026)
Mean mig. rate high qual. district	-0.018	-0.014
	(0.046)	(0.032)
Perc. parents district	0.067	0.074
	(0.035)	(0.029)
Share High qualified in district	0.098	0.098
	(0.035)	(0.034)
No qualification	0.165	0.161
Some qualification	0.835	0.839
High qualification	0	0
Number of observations	27667	22150
Std. Dev. in Parentheses		

Table 4: Regression results: Dependent Variable: yearly wage growth of unskilled and skilled with a-level degree 1993-1995 and 1995-1997

	(1) OLS 93-95	(2) IV 93-95	(3) OLS 95-97	(4) IV 95-97
age	-0.031**	-0.031**	-0.023**	-0.023**
	(48.58)	(48.29)	(43.89)	(43.70)
age_2	0.000**	0.000**	0.000**	0.000**
	(43.14)	(42.88)	(39.17)	(38.98)
Skilled workers	-0.084**	-0.084**	-0.060**	-0.061**
	(31.71)	(31.49)	(26.36)	(26.36)
Clerks and foremen	-0.056**	-0.055**	-0.042**	-0.042**
	(19.82)	(19.56)	(17.66)	(17.60)
Tradable Goods	0.056**	0.056**	0.025**	0.025**
	(15.73)	(15.70)	(7.37)	(7.41)
Nontradable Goods	0.036**	0.039**	0.016**	0.016**
	(8.18)	(8.63)	(4.61)	(4.60)
Marital status	-0.000	0.000	0.003	0.003
	(0.17)	(0.07)	(1.64)	(1.55)
Male	0.012**	0.012**	-0.002	-0.002
	(5.62)	(5.60)	(1.00)	(0.90)
Urbanised area	-0.006**	-0.012**	-0.005**	-0.002
	(3.10)	(4.90)	(2.77)	(1.19)
Rural area	-0.006*	-0.008**	-0.002	0.003
	(2.45)	(3.03)	(1.12)	(0.82)
Unempl. rate Land/Educ/Branch	0.412**	0.459**	0.325**	0.324**
	(10.21)	(10.93)	(9.29)	(9.22)
Mean mig. rate high qual. district	-0.028	-0.337**	-0.038	-0.310*
	(1.45)	(4.56)	(1.64)	(2.31)
Share High qualified in district	0.066*	0.125**	0.048*	0.134**
	(2.45)	(4.11)	(1.97)	(2.77)
Constant	0.684**	0.666**	0.506**	0.491**
	(45.45)	(42.54)	(39.10)	(33.27)
Observations	27664	27664	22150	22150
Adjusted R-squared	0.21	0.21	0.21	0.20

Absolute value of t statistics in parentheses
+ significant at 10%; * significant at 5%; ** significant at 1%
Omitted Categories: Unskilled Workers, Agriculture and Mining, Agglomeration area
Instrument in IV is the share of parents in the district

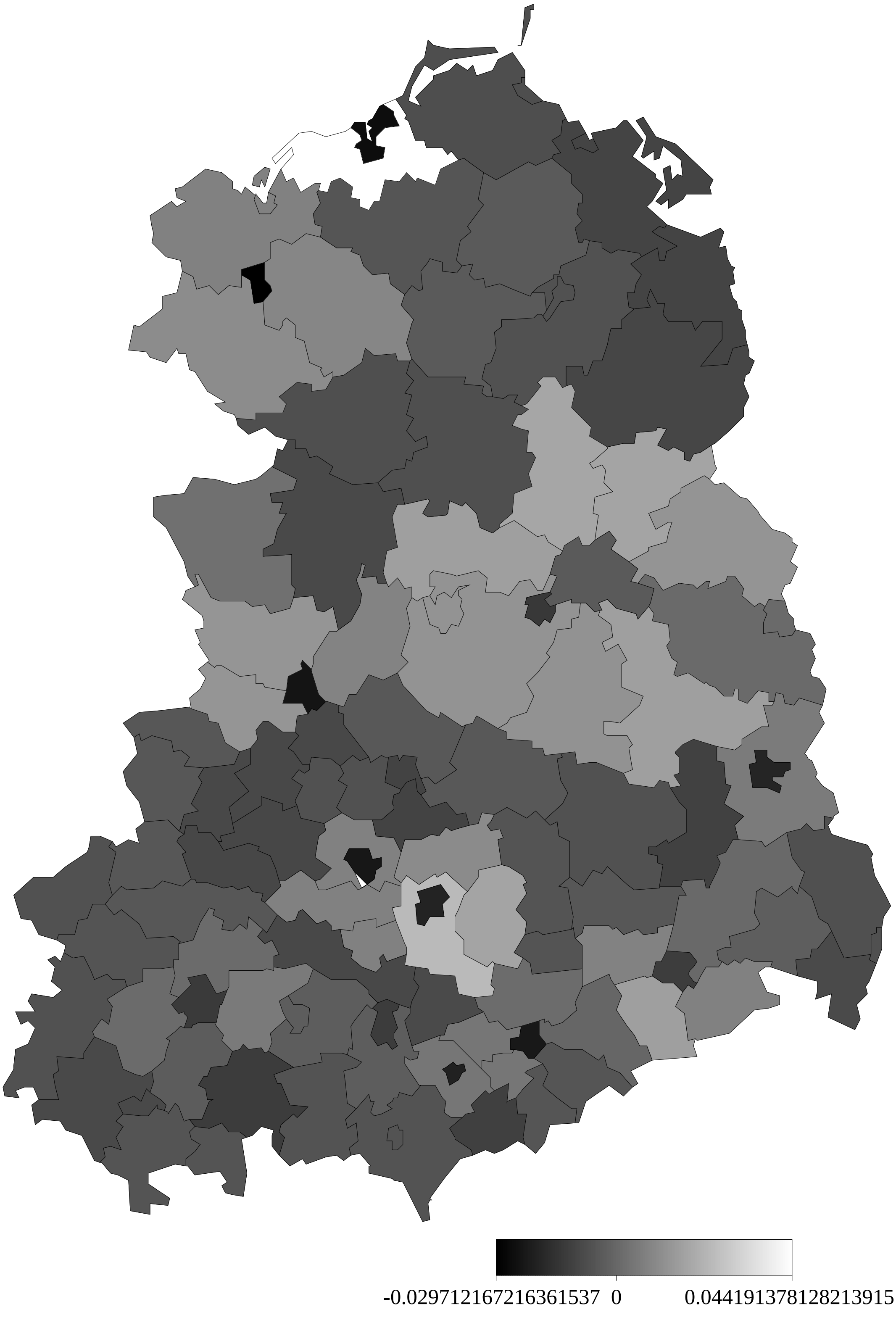

Figure 2: Regional population growth 1993-1997

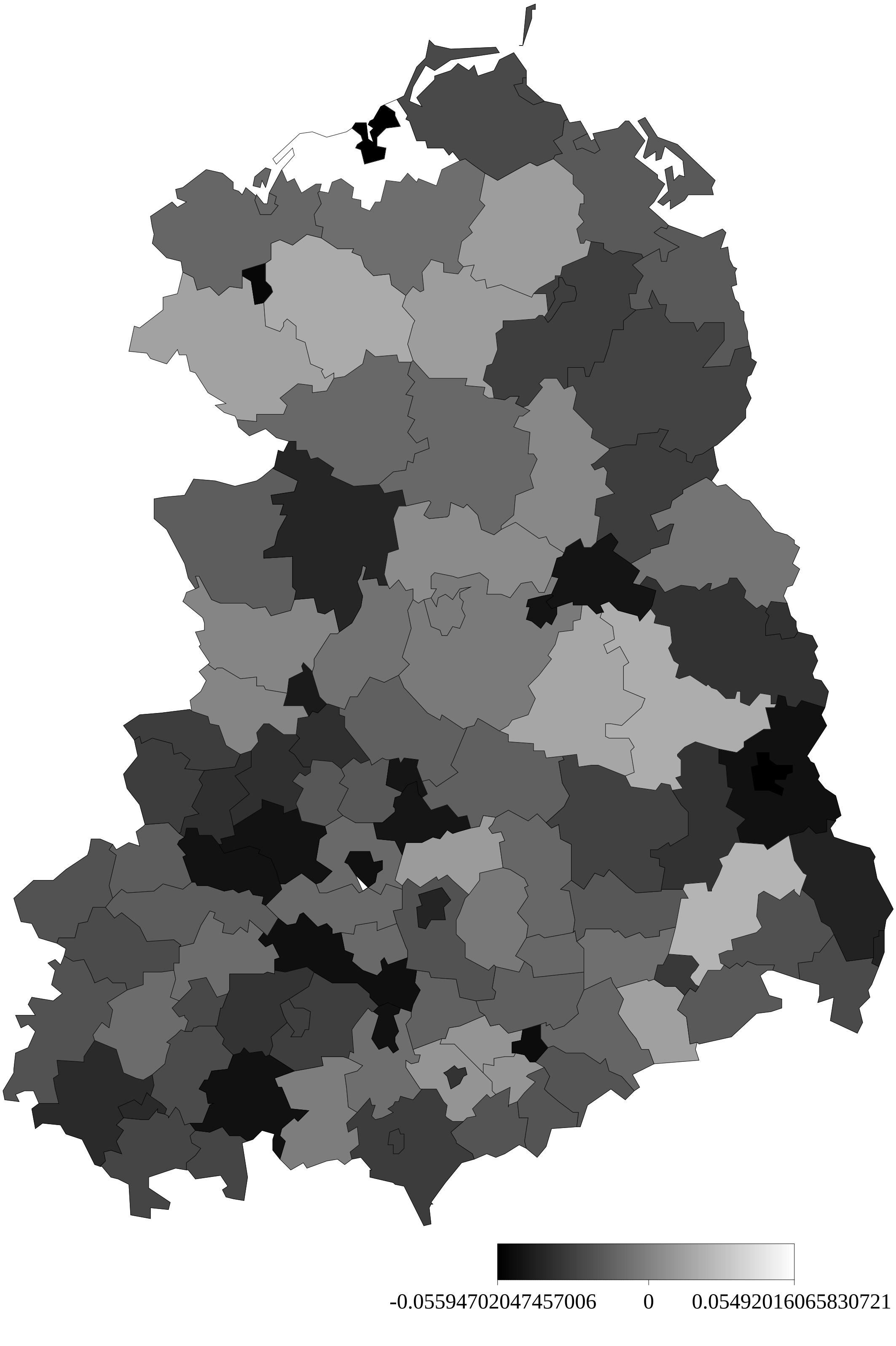

Figure 3: Regional employment growth 1993-1997

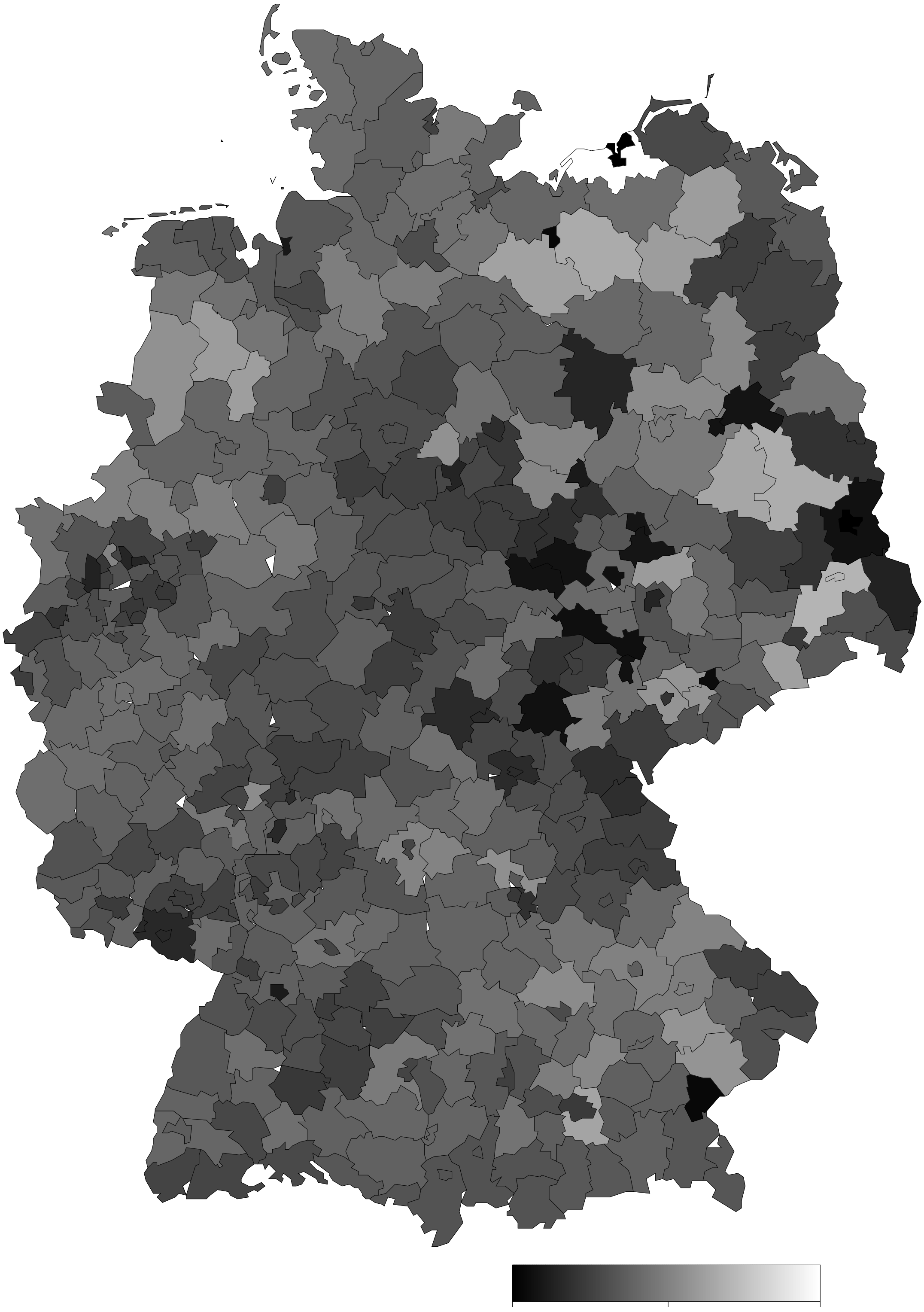

Figure 4: Regional employment growth Germany 1993-1997

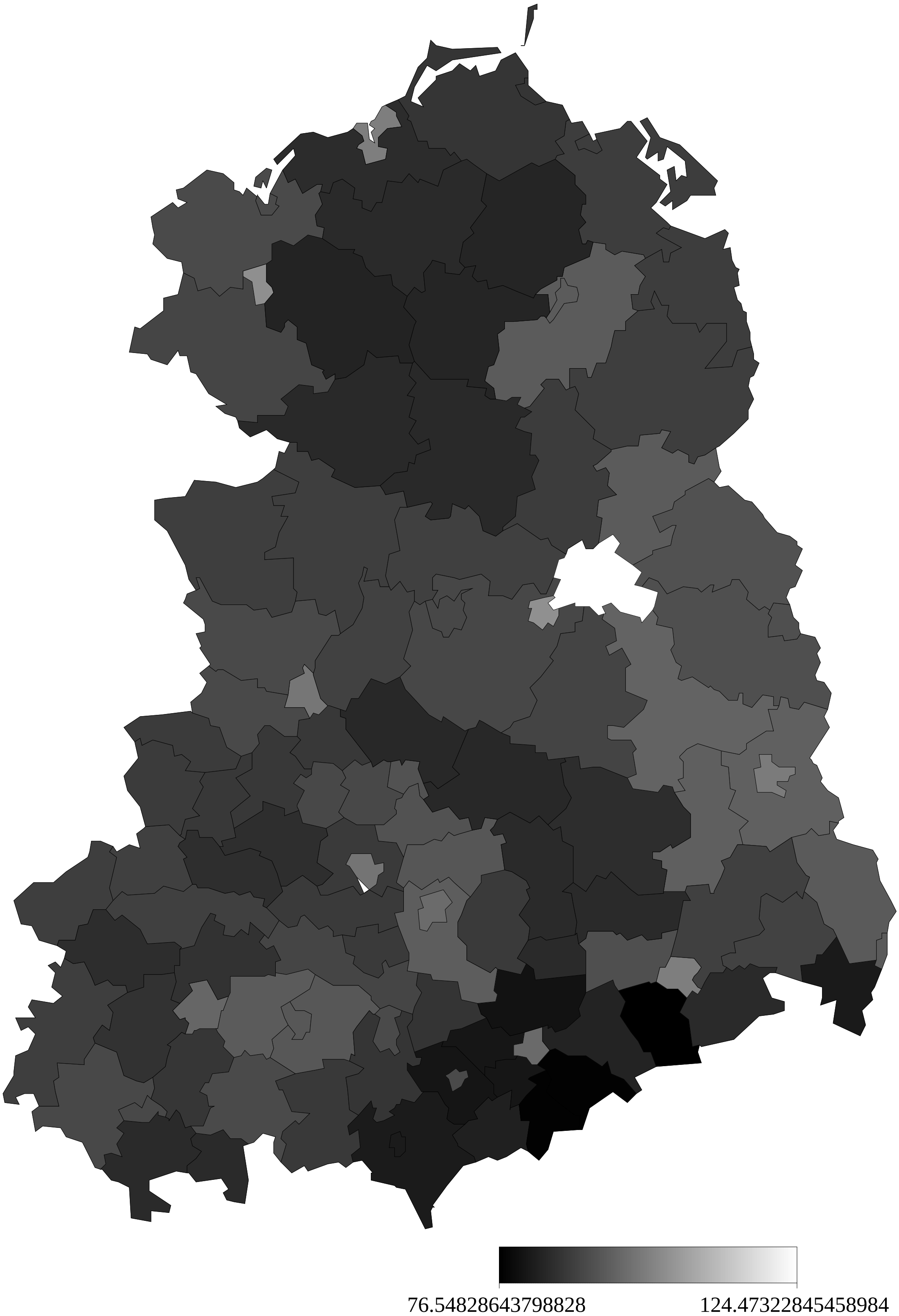

Figure 5: Regional wage levels in 1993

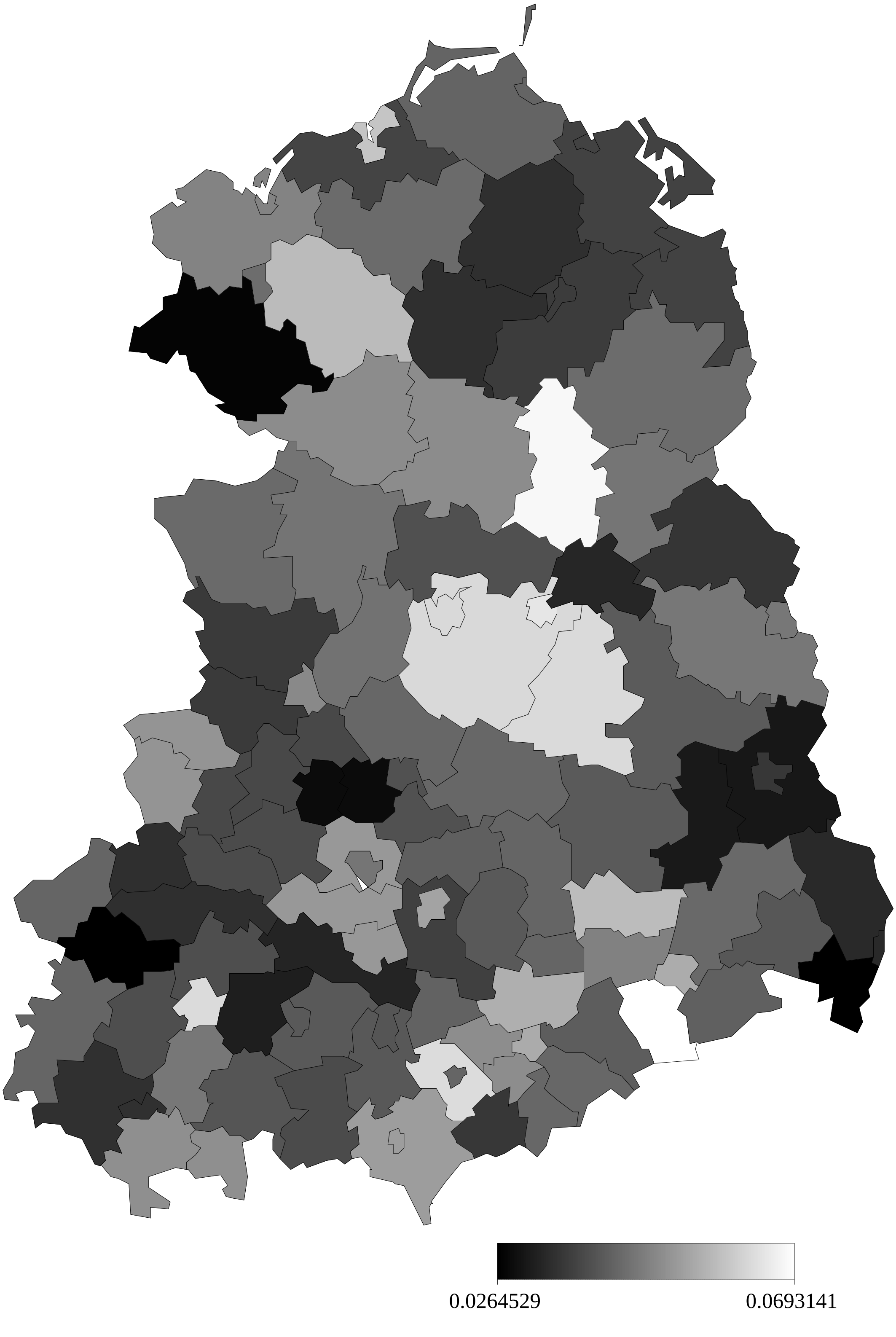

Figure 6: Regional wage growth 1993-1997

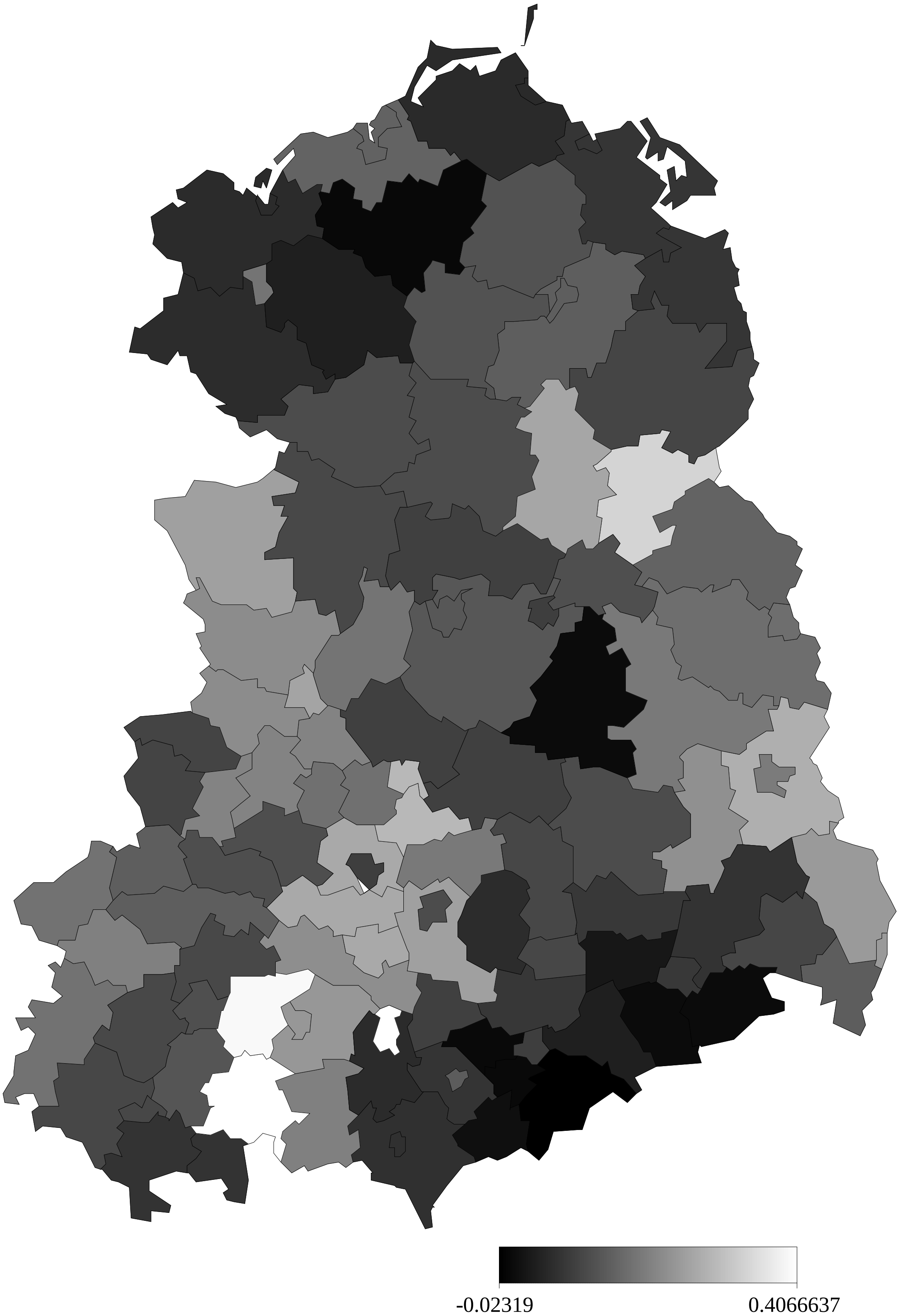

Figure 7: Regional unemployment growth 1993-1997

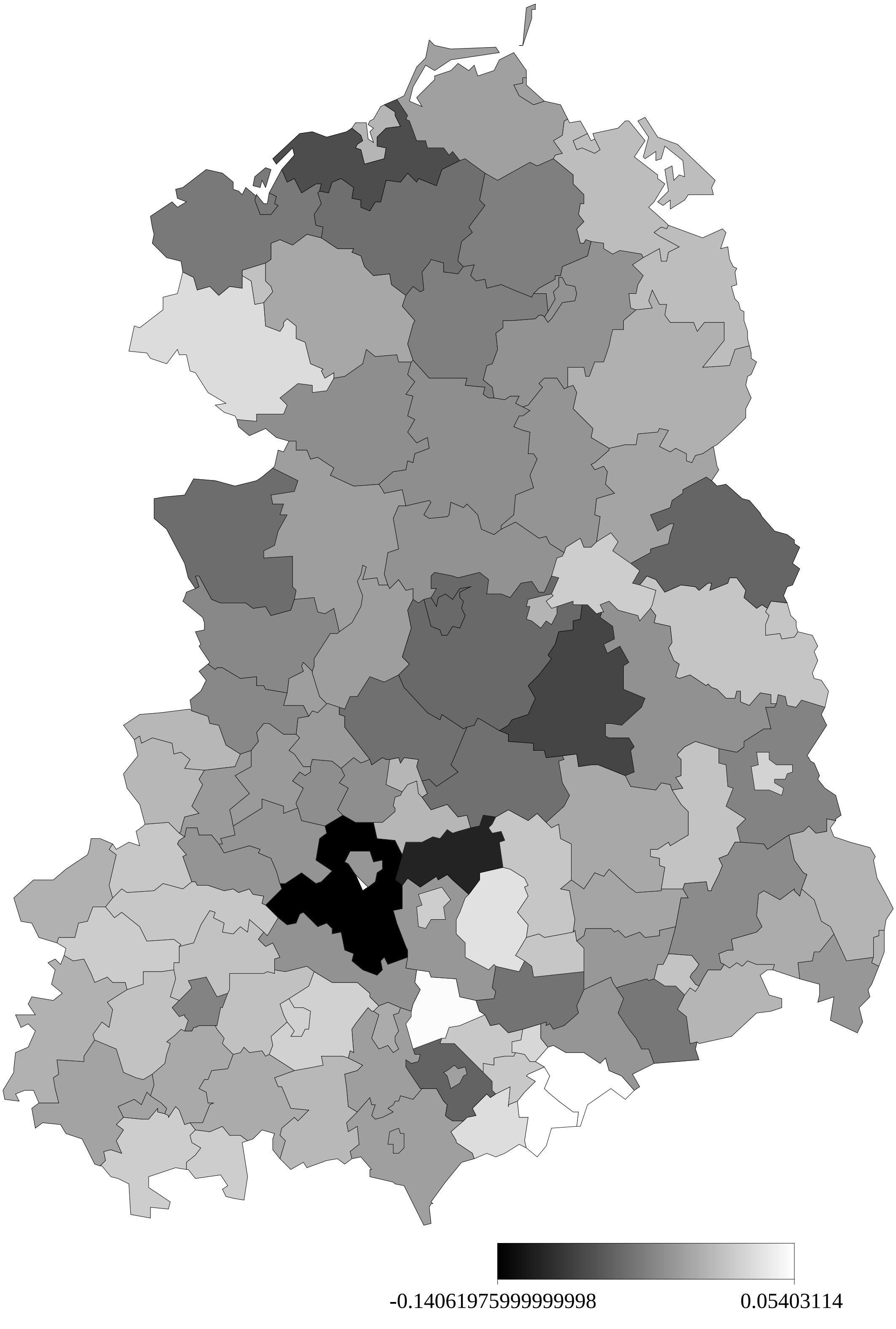

Figure 8: Mean regional migration rates of high-skilled 1993-1997

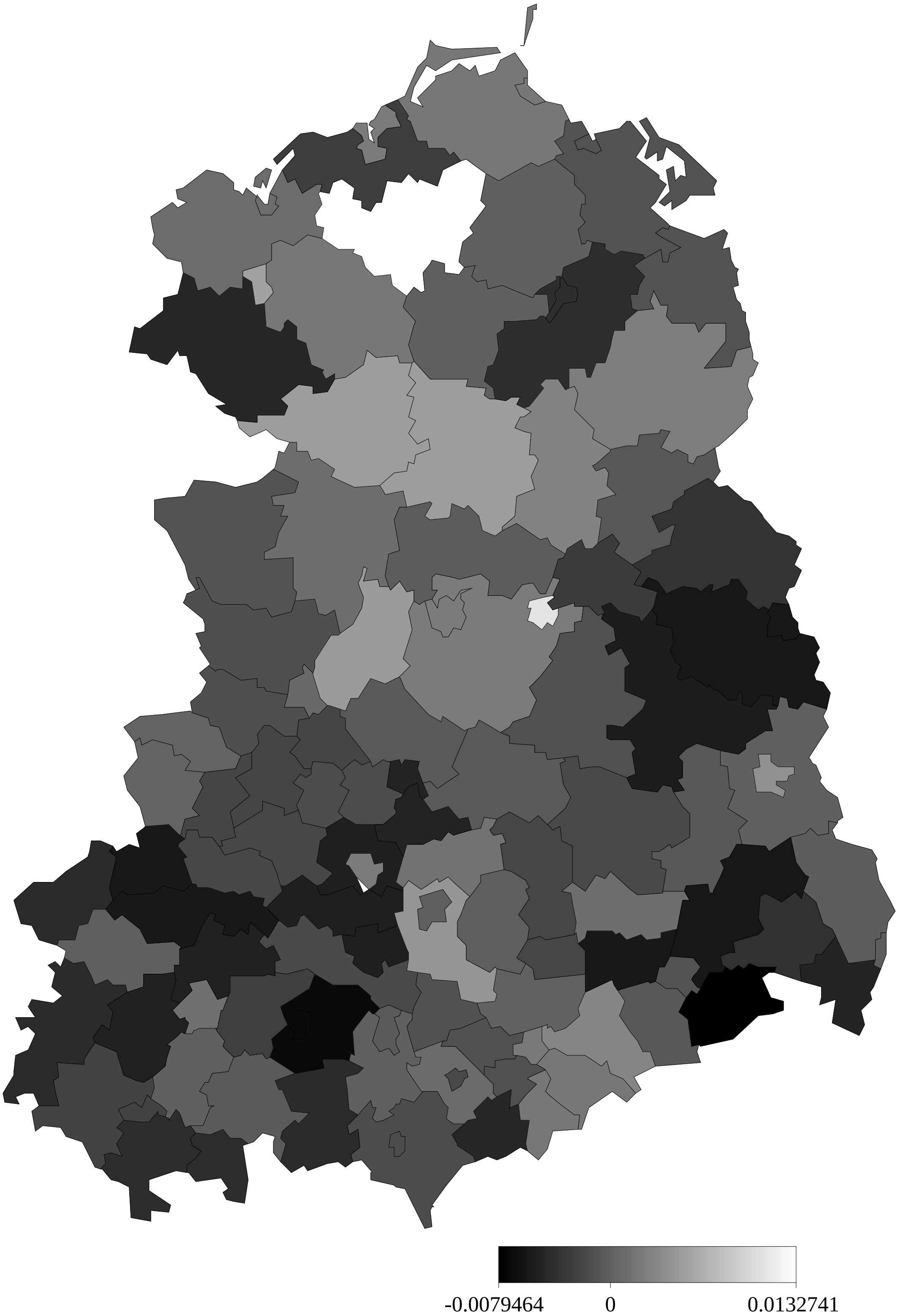

Figure 9: Change in the share of highly educated workers 1993-1997

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