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E-mail address: cornelius.hirsch@wifo.ac.at, harald.oberhofer@wifo.ac.at

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Bilateral Trade Agreements and Trade Distortions in Agricultural Markets*

Cornelius Hirsch

Austrian Institute of Economic Research
Cornelius.Hirsch@wifo.ac.at

Harald Oberhofer

Vienna University of Economics and Business
Austrian Institute of Economic Research
Harald.Oberhofer@wu.ac.at

Abstract

Agricultural support levels are at a crossroad with reduced distortions in OECD countries and increasing support for agricultural producers in emerging economies over the last decades. This paper studies the determinants of distortions in the agricultural markets by putting a specific focus on the role of trade policy. Applying various different dynamic panel data estimators and explicitly accounting for potential endogeneity of trade policy agreements, we find that an increase in the number of bilateral free trade agreements exhibits significant short- and long-run distortion reducing effects. By contrast, WTO's Uruguay Agreement on Agriculture has not been able to systematically contribute to a reduction in agriculture trade distortions. From a policy point of view our findings thus point to a lack of effectiveness of multilateral trade negotiations.

JEL classification: C23; C26; F13; F14.

Keywords: Agricultural distortions; WTO; bilateral trade agreements, panel data.

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1 Introduction

Studying the economic rationale for governmental interventions in markets and understanding its outcomes constitute one of the main and long-lasting questions addressed in economics. Thereby, researchers have put a specific emphasis on the food and agriculture sector which has been subjected throughout history and all over the world to the most extensive and far-reaching political steering and intervention attempts. From the historically most famous protectionism in Britain in the 19th century - the Corn Laws - to the costly Common Agriculture Policy (CAP) in the EU or the outstandingly high export taxes in South America, there are numerous cases to study the causes and consequences of such extensive market interventions. In recent economic research the question on the main determinants of market interventions has been mainly studied in the field of political economy that has come a long way from niche player to a suitable analysis tool for economic policy reforms. Thus for example the World Bank conducted a yearlong research project on agricultural distortions with a key focus on political economy analyses and triggered a new wave of publications surveyed in e.g., Anderson, Rausser, and Swinnen (2013). This research project provides new and interesting findings on recent trends regarding market distortions in agricultural markets, motivating further research especially in light of changing international trade relationships due to multilateral and bilateral trade policy initiatives.

Over the last two decades a heated debate emerged on how to most effectively organize trade negotiations. After the conclusion of the Uruguay Round Agreement and the founding of the WTO, multilateral trade talks have become less successful and different proposals for further rounds of trade liberalization have not made substantial progress, did not achieve unanimous assent from all members and thus could not be approved. As a consequence, we observe a tremendous rise in the number of signed bilateral (and plurilateral) trade agreements leading to the formation of regional trade blocs (see, e.g., Bagwell, Bown, and Staiger 2016). From 1991 to 2010 to number of bilateral trade agreements has been steadily rising from 51 to more than 400 (Dür, Baccini, and Elsig 2014). As a result, trade negations are nowadays carried-out simultaneously at the multilateral as well as the bilateral level and countries might use just one of both options in case they believe they will be more successful by concentrating on a specific type of trade negotiations. In line with this development, a comprehensive analysis of the impact of trade policies on market distortions in agricultural markets needs to simultaneously account for both multilateralism as well as bilateralism. While the impact of the formation of the WTO (ignoring bilateral trade agreements) on agricultural markets has been already studied in the literature (see, e.g., Swinnen, Olper, and Vandemoortele 2012), a simultaneous analysis of WTO membership and bilateral trade negotiations is still missing.

With the shifting balance of the global economy in the last half decade in terms of dynamic GDP growth in emerging countries and the formation of new regional trade blocks, the overall picture on policy support granted to the agricultural sector has changed significantly. Numerous lower income countries have switched from taxing agricultural production to applying protectionist measures, while richer countries reduced their level of most distorting measures and shifted their policy support towards decoupled payments to farmers (Swinnen et al. 2012). Based on previous experiences the rising support levels in emerging markets and other fast growing lower income countries can be expected to have a significant impact on international agricultural markets as their overall economic relevance is will continue to grow.

Therefore it is of great interest to understand the impact of macroeconomic factors and other international trade issues on the development of agricultural support measures. In particular, the aim of the paper is to empirically investigate the impact of trade negotiations on market distortions in (international) agricultural markets and to disentangle the respective contributions of multilateral versus bilateral trade talks. This paper builds to a certain extent on Swinnen et al. (2012) in analyzing the level of distortions but extends the analysis to different types of trade agreements and also utilizes an updated and enlarged dataset, in which all necessary data are available for a time span capturing the years from 1980 to 2011 and 76 countries. Accordingly, we apply a dynamic panel data framework which allows to identify the ceteris paribus effects for both different types of trade agreements and enables us to calculate both the short- and long-run effects of trade policy on market distortions in agricultural markets. In order to account for the persistence in market distortions over time, World Bank's nominal assistance coefficient (discussed in detail in Anderson et al. 2013) is regressed on its one year lagged value and on bunch of control variables identified as crucial determinants for market distortions in the previous literature. On top of that, the specification accounts for a country's WTO membership as a measure for multilateral trade agreements together with its number of bilateral trade agreements in force obtained from Dür et al. (2014). In the empirical analysis we consider different dynamic panel data estimators also explicitly addressing the potential endogeneity of the number of bilateral trade agreements signed.

In line with previous literature, our estimates indicate a non-linear relationship between market distortions and the average income level (measured in terms of GDP per capita). Accordingly, market distortions are increasing with income but at a decreasing rate. Furthermore, countries which exhibit a positive agricultural trade balance tend to shwo less positive or in some cases even negative price distortions. Furthermore, agricultural market distortions are characterized by non-negligible persistence over time. With regard to the variables of most interest, a country's WTO membership does not systematically affect a country's distortions in agricultural markets while an increase in the number of signed bilateral trade agreements significantly reduces agricultural market distortions both in the short- and long-run. In particular and based on the preferred econometric specification and estimator, a one standard deviation increase in the number of bilateral agreements (i.e., 22.3 signed agreements) decreases the nominal assistance coefficient by -0.089 and -0.309 in the short- and long-run, respectively. From a policy point of view, our findings suggest that multilateral trade agreements as negotiated within the WTO-framework lack effectiveness in terms of contributing to reducing distortions in international agricultural markets. By contrast, bilateralism and the formation of regional trade blocs seem to be more effective in reducing market distortions in agricultural markets.

The reminder of the paper is structured as follows: In Section 2 we discuss the main political economy arguments identified in the literature on market distortions in agricultural markets and describe the main recent trends and developments in agricultural policies taking place all over the world. Section 3 provides and in depth discussion on the utilized data, the empirical specification applied and highlights the most important econometric issues involved for the identification of short- and long-run effects of trade policy for distortions in agricultural markets. In Section 4 we present our estimation results and provide an extensive sensitivity analysis, while Section 5 offers some concluding remarks.

2 The Political Economy of Agricultural Policy

This section offers a discussion on the most recent trends and developments in agricultural markets distortions and provides a brief overview on the existing empirical literature studying the main determinants for market distortions. Furthermore, we also briefly present data on bilateral trade agreements highlighting their increasing relevance for international trade policy making. This section thus motives our main research question and embeds the empirical framework proposed into the relevant literature.

2.1 Recent Trends and Developments in Agricultural Policy

In the history of agricultural economics and international trade one event is unanimously referred to as a milestone in trade liberalization – the repeal of the Corn Laws in the 19th century. With the end of the Napoleon wars in 1815, grain prices fell after the agriculturalists had enjoyed a period of high prices during the war times. These generally lower prices in combination with superior yields in France and the Netherlands compared to those in England led to new legislation in England to keep out cheap corn. In contrast to the preceding corn laws, which existed for many decades, the 1815 corn laws were "definitely protective" (Adams 2013; Schonhardt-Bailey 2006). The influential landowners organized in the Conservative Party benefited from the protectionism and it took until 1846 that, after strong opposition from a growing industrial middle class, the Corn Laws were repealed. Economic historians see the repeal of the Corn Laws as a major boost for British free trade and, in consequence, for the industrialization of Europe. The repeal of the Corn Laws represents without a doubt on of the biggest shifts in agricultural support policies in history and remains a much debated topic for political economists and agricultural economists alike. The history of the Corn Laws does, however, already highlight two key aspects influencing agricultural price policies: The agricultural trade balance and the overall economic development played an important role both for introducing and repealing the protective measures.

Great Britain's commitment to free trade - not only in agriculture - lasted longer than in most other western regions and only with the economic depressions in the 1920s the country returned to import duties (UK Parliament 2016). After the Second World War western countries returned to free trade policies, most notably marked by the initiation of negotiations on the General Agreement on Tariffs and Trade (GATT). Accordingly, the second event that economists commonly refer to when it comes to trade liberalization in agriculture is the GATT's Uruguay Round Agreement on Agriculture (URAA) in 1994. As a consequence of the multilateral trade negotiations, the second half of the 20th century has seen several striking developments in regard of agricultural distortions. Distortions to agricultural markets through policy measures have undergone significant changes with very different trends occurring in different geographical regions in the world.

The last three decades have seen several attempts of analyzing the degree of agricultural distortions and the drivers for reforms in different countries all over the world. All empirical publications are based on either one of two (across countries comparable) data sources on this matter. Since 1987 the OECD offers a yearly monitoring on agricultural support and an online database with several indicators for agricultural distortion, covering 23 OECD and Non-OECD countries.[1] In addition, in 2010 an extensive World Bank research project collected and estimated agricultural distortion data and updated this database in 2013.[2]

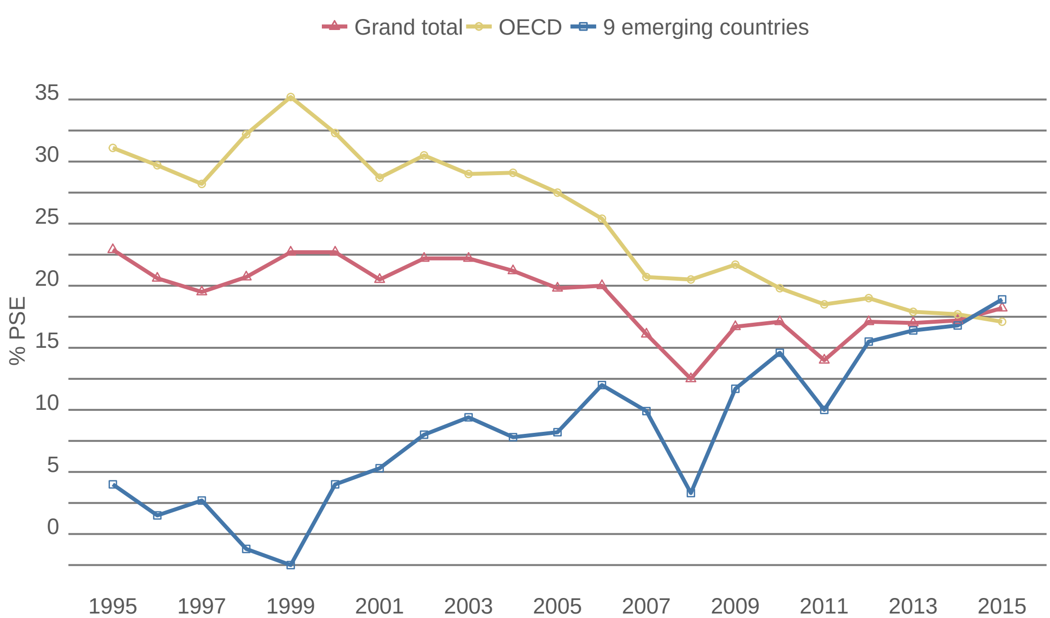
Figure 1 and 2 provide a descriptive overview on the developments of the different measures for agricultural distortions over time and by also distinguishing across different country groups. The OECD's Producer Support Estimate (PSE) measure is defined as total of all *"policy transfers to agricultural producers, measured at the farm gate and expressed as a share of gross farm receipts"* (OECD 2016).

[1]The OECD database can be accessed online via www.oecd.org/tad/agricultural-policies/producerandconsumersupportestimatesdatabase.htm.

[2]The World Bank database is available at www.worldbank.org/agdistortions (Anderson and Nelgen 2013).

According to Anderson (2009), *"[T]he nominal rate of assistance (NRA) is defined as the percentage by which government policies have raised gross returns to farmers above what they would be without the government's intervention (or lowered them, if NRA < 0)"*. The former measure for market distortions in agricultural markets is thus directly linked to all income-related support measures granted to farmers, while the NRA more systematically captures the price effects of any policy support. As indicated in Figure 1, the overall level of support for agriculture follows a downward trend over the time period spanning the years from 1995 to 2015 captured by the OECD's PSE database for agricultural policies and support. However, this observation should not detract from the fact that the developments are heterogeneous across advanced and emerging market economies. While in the former support has been declining as a percentage of gross farm receipts, total agriculture support is increasing in the latter group of countries. Given the opposite starting position in 1995, this leads to a convergence in agricultural markets support and, since 2014, the non-OECD countries offer more generous policy support for farmers as compared to the included OECD-member states.

Figure 1: Evolution of Producer Support Estimate (PSE) , 1995 to 2015 (percentage of gross farm receipts).

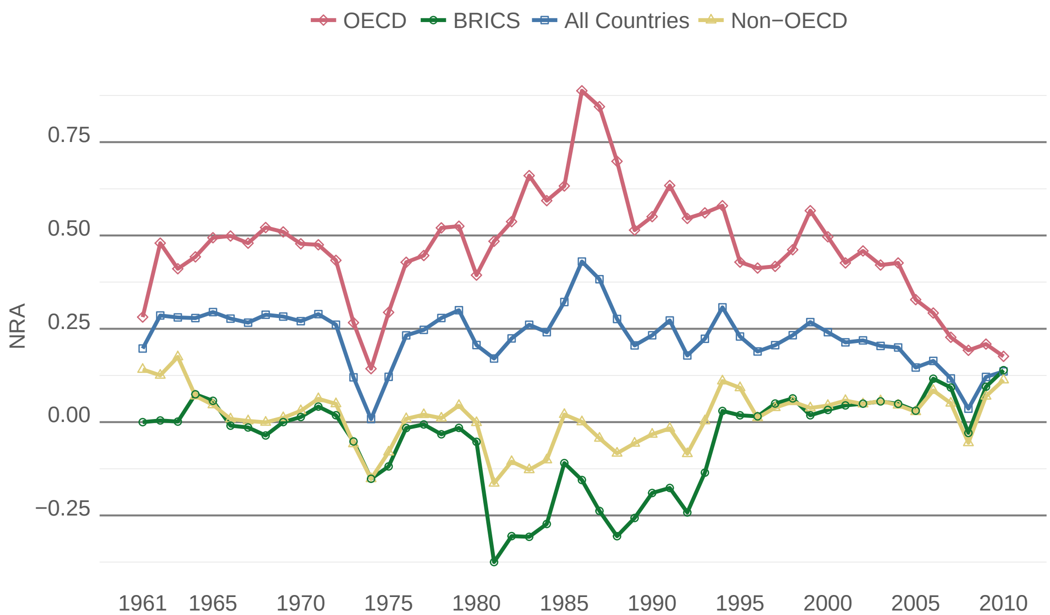


Notes: Source: OECD (2016). The OECD total does not include the non-OECD EU Member States. The Czech Republic, Estonia, Hungary, Poland, the Slovak Republic and Slovenia are included in the OECD total for all years and in the EU from 2004. The emerging economies are Brazil, China, Colombia, Indonesia, Kazakhstan, Russia, South Africa,Ukraine and Viet Nam. Viet Nam is included from 2000 onwards.

The NRA index provided by a World Bank project is available for a longer time period spanning the years from 1961 to 2011 and its development over time can be based on a much broader country sample consisting of 81 countries.[3] Furthermore, the NRA is expressed as percentage of the undistorted price which is not the case for OECD's PSE measure which considerably simplifies its interpretation (Anderson 2009). Unfortunately, the dataset has only been updated until 2011 which does not allow to account for the most recent developments. However, a visual inspection of Figure 2 indicates that the NRA captures very similar dynamics as the PSE indicator. In particular, in the OECD economies distortions in agricultural markets have been at a rise until the late 1980s while policy support for

[3]Due to data limitations regarding the control variables necessary for the econometric analysis discussed in Sections 3 and 4, Côte d'Ivoire, Island, Malta, Sudan and Taiwan can only be included in this descriptive section.

Figure 2: Evolution of Nominal Rate of Assistance (NRA) in OECD, non-OECD and BRICS countries, 1961 to 2011.



Notes: Source: Anderson and Nelgen (2013). NRA: Nominal Rate of Assistance is an index for distortions to agricultural prices. BRICS countries include Brazil, Russia, India, China and South Africa. The displayed values are weighted averages of the countries' NRA by their value of agricultural production.

farmers substantially decreased afterwards. In the average OECD economy (based on agricultural GDP share-weighted country average), in 2011 the market distortion in terms of a percentage increase relative to the undistorted price amounts to only less than one-third of the relative distortion observed in 1986. Furthermore, Figure 2 also indicates that in lower middle income countries and especially the BRICS economies, agricultural support levels have been growing over the last three decades. During the 1980s and until the mid-1990s, these economies have taxed agricultural outputs above the standard value-added tax level for other goods which translated into a negative market distortion and lower prices for producers. Since 1995, however, the group of lower middle income countries and the BRICS economies have started to (heavily) subsidize agricultural production and until 2011 almost reached the level of market distortions (in terms of above undistorted market prizes) of the OECD economies.

The process of agricultural policy setting seems to follow to certain extent a pattern which is consistent with the descriptive evidence provided in Figures 1 and 2. Three observed phases along the process of economic development have been identified: Countries at the low income levels could regard their large agriculture sectors as source for governmental revenues and tax agriculture to extract resources from the sector and facilitate non-agricultural developments. With accelerated economic growth, countries often tend to provide net support to the agriculture population as their incomes do not keep pace with non-agricultural sectors and the reallocation of their labor force faces difficulties. Governments therefore tend to transfer income to the agricultural sector to mitigate these disparities. The third phase taking place at a higher economic development stage is characterized by new policy objectives gaining importance in the respective countries. In order to comply with the new topics such as environmental concerns and sustainability requirements, competitiveness of the sector and stabilizing farm incomes, governments tend to shift their focus towards less market price distorting measures but rather payments which are subject to specific conditions (OECD 2013).

As a consequence and for an accurate interpretation of the overall developments in agricultural policy, it is important to distinguish between distortions and support. As among others Swinnen et al. (2012) argue that while the total level of support to farmers has not changed much in OECD countries, the most distortive measures indeed have been abandoned. This is due to a shift towards policies such as direct payments which have less or no effect on prices faced by farmers and are not considered as market distorting (Swinnen et al. 2012). This paper is interested in analyzing agricultural distortions to markets and, therefore, relies on the NRA as the preferred indicator which measures price differentials induced by agricultural policy measures.

2.2 A Brief Review of the Literature

There is little dispute in the literature over the observed patterns of agricultural support and the simultaneous occurrence of changes in macroeconomic conditions and agricultural support. However, causality is more complex and of high interest if one wants to elaborate on possible future scenarios for agricultural policies. This is especially relevant given the level of support in emerging market economies is now exceeding the ones provided by the OECD countries which could possibly lead to a turnaround in the gross total support of all countries. In return this is expected to have a significant impact on world agricultural commodity markets. Academics and policy advisers have been puzzled by the determination of trends in agricultural support and how their developments seem to be resisting any advocacy of greater market-orientation. Political economy arguments have been heavily involved in analyzing and explaining the determining factors of changes in agricultural support policies. From the numerous empirical studies conducted over the last decades three main patterns based on empirical evidence can be identified (see Swinnen et al. 2012, for a survey): First, the development pattern refers to the already mentioned shift in policies from taxing the agricultural sector to subsidizing and protecting the same with increasing rates, as countries move up the scale of economic development. Second, the anti-trade pattern is based on the observation that import competing sectors generally witness more protection than exporting sectors. In addition, net-exporting economies tend to generally provide lower levels of agricultural support. Third, the anti-comparative advantage pattern refers to the negative correlation between comparative advantage in and support for an agricultural production sector.

The macroeconomic and structural indicators, which the observable patterns refer to, correlate with political economy determinants of agriculture support. The relationship between various structural variables is well explored, such as the decreasing share of agriculture in the economy with a growing economy but the connection to agriculture support is less thoroughly examined. Yet, the World Bank focused on the relationship between the development of distortions to agricultural incentives and political economy reasoning in a large research program carried out during the years from 2006 to 2009, providing a detailed summary of the main findings in e.g., Anderson (2010).

The theoretical literature on endogenous policy determination suggests that agricultural protection tends to rise with increasing GDP and declining share of agriculture in the economy and when consumers spend a lower share of their budget on food (Brooks 1996). With economic development and growing GDP, per capita incomes increase and, following Engel's law, a smaller share of a household's budget is spent on food (Engel 1857). In combination with Downs' claim of consumers' rational ignorance increasing support faces less resistance. Downs (1957) called consumers and taxpayers rationally ignorant when they pay less attention to government policies where the costs of becoming fully informed about the policy effects outweigh the potential individual benefits stemming from alternative policies. Since the costs of agricultural programs affect consumers and taxpayers less as they spend a smaller proportion of their increasing incomes on food they have fewer incentives to exert countervailing pressure.

Moreover, as the economy grows and the share of agriculture in total employment declines, implying that the size of the potential interest group of farmers shrinks, their collective action becomes more effective. Olson (1965) famously stated that free riders present the main obstacle to effective collective action such as lobbying for more policy support. With diminishing numbers, groups find it easier to overcome this free-rider problem because the incentive to actually free ride on group activity increases with the size of the group and therefore the coordination and enforcement costs increase. Buchanan and Tullock (1962) argued already in an earlier contribution that collective decision making will be more effective in smaller groups, because of lower participation and organizational costs involved in the policy process.

Another observation of past developments is that, as the country's economy develops, incomes in the non-agricultural economy typically grow at a faster rate than in the agriculture sector. The ongoing competitive process forces farmers, who cannot match cost reduction, out of the market unless they look for non-market sources of income such as e.g., governmental support. Clearly, farmers who may be forced to leave the sector have a high incentive to lobby for more support (Brooks 1996). Similar reasoning could explain the observed greater policy support in sectors with a comparative disadvantage. The incentive for governments to exchange transfers for political support is increasing with declining relative incomes in the farming sector. This relationship has been first described formally by De Gorter and Tsur (1991). Therefore, the anti-comparative advantage pattern is often also referred to as a relative-income pattern (Swinnen 2010).

The last macroeconomic variable to discuss in the context of endogenous policy determination is the agricultural trade balance. The literature predicts that agricultural exports will be subsidized less (or taxed more) as compared to agricultural imports. The costs of distorting transfers increase with an increasing trade surplus, because both the dead-weight loss of the distortion itself and the actual transfer costs increase with higher subsidization (Brooks 1996). The motive of intervening in trade in a way to lower domestic prices is again related to the share of consumers expenditures on food and the effect on government border tax revenues (Swinnen 2010). Anderson (2010) points to another reason regarding the observed anti-trade patterns in developing countries over the past: *"Part of the anti-trade bias in developing countries was the result of government intervention in the domestic market for foreign currency. The most common arrangement was a dual exchange rate, whereby exporters had to sell part or all of their foreign currency to the government at a low price. This effectively taxed and thus discouraged production of exportables. At the same time it created an artificial shortage of foreign currency so that potential importers bid up its purchase price, which had the same effect as an import tax and thus encouraged import-competing production"*. Yet another argument in addition to the already mentioned costs associated with supporting exports could be that countries tend to face stronger retaliation from other countries as more subsidized exports flood the international market, whilst support for imports may cause no policy response from the trading partners.

One could argue that the most relevant factor affecting agricultural policy is constituted by international trade agreements. The members of the former GATT and now WTO have been reluctant to agree on multilateral agreements especially in agriculture. Nevertheless, when they did, the impact on support policies was profound across countries. Over recent decades, from the Uruguay round (URAA), the establishment of the WTO and the Doha round up to the Bali package and the latest Nairobi package the stepwise deepening of multilateral trade agreements has been shaping agriculture support policies. As already discussed in the introduction, the URAA is often referred to as one of the key events of liberalization in world trade and as a consequence thereof it is subject to extensive scientific discussions. Surprisingly, there is little econometric work on this issue, at least not to the extent as it attracts interest in the discussions. The general consensus of experts seems to be that while the URAA may have constrained the growth of agricultural protection it has done little to reduce it, at least in the countries that were GATT members during the negotiations (Anania, Bohman, Carter, and McCalla 2004).

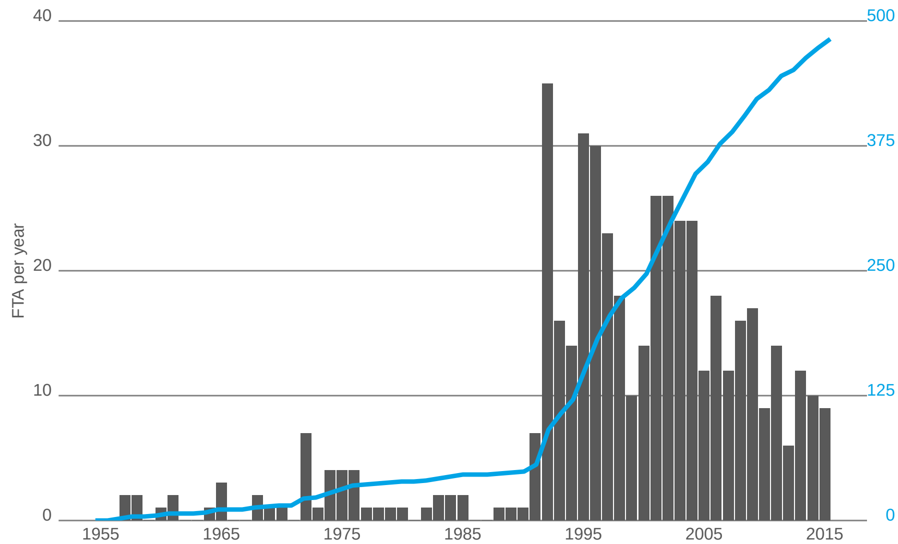
A potential explanation for why the URAA has had no or limited impact on market distortion in agricultural markets can be found in the tariff regulation agreed under WTO rules. For developing economies the tariff rates are fixed well above the actual applied tariffs at the time the negotiations took place. Thus, for example China bound its tariffs at an average level of 16.5 percent while its respective NRA at the time it joined the WTO only amounted 7.3 percent. This has been similarly true for several other developing and emerging economies (Anderson 2010).

Furthermore, there are several obstacles for adequately measuring and quantifying the impact the URRA agreement. One is that they may be anticipated and thus the policy change actually happened prior to the official signature of the agreement. Furthermore, the starting conditions in terms of a country's embeddedness in (already) existing trade agreements and the extent of agricultural markets liberalization differed across countries which complicates the identification of any WTO effects. *"For example, among the 14 transition countries the impact of the GATT/WTO on their agricultural policies differs strongly depending on whether they were part of the GATT before 1995 or not"* (Swinnen 2010). Nevertheless, trade talks certainly do play a role, as do transparent information and policy advice, provided for example by the OECD with its public and regularly updated database on agricultural support. However, recent empirical contributions such as e.g., Swinnen et al. (2012) could not find a significant market distortion reducing effect of the URAA using OECD data. Cadot, Olarreaga, and Tschopp (2010) find a weakly significant effect of the URAA in terms of reducing the volatility of agricultural support over time.

Besides the multilateral trade talks within the WTO context, bilateral free trade agreements have gained importance in international trade relations and are expected to influence both the degree of market access and distortion to agricultural markets. The subsequent analysis offered in this paper utilizes the Design of Trade Agreements Database (DESTA) which, as of June 2016, contains more than 800 agreements. This database, thus, most comprehensively captures trade agreements and, in contrast to most other trade agreements listings, provides additional information on the depth of the respective agreements (Dür et al. 2014). According to DESTA the number of bilateral trade agreements has increased significantly over the last 25 years and according to Dür et al. (2014) the signed agreements also became deeper in terms of their scope and how comprehensively they regulate trade relationships. Figure 3 illustrates this point by plotting the number of annually signed bilateral free trade agreements together with the cumulative number of such agreements in force. Until the early 1990s, bilateral free trade agreements have been negotiated only very infrequently and, as a consequence, a total of 100 agreements in force has been reached only in 1993. Starting with the 1990s the number of annually signed bilateral free trade agreements almost exploded, reaching a maximum of 35 signed bilateral agreements in 1992. Furthermore, this trend in the popularity of signing bilateral free trade agreements continues until today amounting to an impressive number of almost 500 cumulative such agreements being in force in 2015.

Whether regionalism and multilateralism are antagonistic forms of liberalization is not at all established in the literature. One side sees them as substituting international attempts, others call and hope these

Figure 3: Number of Signed Free Trade Agreements, 1955 to 2015).



Notes: Source: Dür et al. (2014). FTA: Bilateral and Plurilateral free trade agreements signed. The blue solid line (right axis) displays cumulative number of free trade agreements in force.

agreements to be a source of traction for multilateral liberalization or even a consequence of the success of multilateral trade agreements (Bagwell et al. 2016; Ethier 1998; Freund 2000; Ornelas 2005). While the verdict on the overall effect of trade agreements on trade and price policies is still out, the notion is that intra-member trade barriers are indeed reduced. Furthermore, free trade agreements may also influence agricultural distortions even in the cases where agriculture is not explicitly a subject of the agreement itself. Intensified bilateral trade relationships might (at least in the long-run) induce market openings in all markets including the one for agricultural products. However, simply interpreting the number of signed trade agreements by a country as a proxy for the economies openness to trade would be oversimplifying. A number of trade agreements do in fact include binding declarations concerning agricultural distortions. Thus, as an example, the North American Free Trade Agreement (NAFTA) provides specific regulation on tariff reduction and tariff quotas for Canadian, Mexican and US agriculture policies (Canada 2016).

Based on the agricultural distortions database the World Bank project conducted a number of empirical analyses concerning the influential factors for market distortions in agriculture. Among them, Masters and Garcia (2010) test the political economy hypotheses focusing on the above discussed development, anti-trade and resource abundance arguments, respectively. The applied OLS-regressions yield significant results for the respective coefficients and confirm the stylized facts obtained from descriptive statistics. In particular, GDP per capita growth increases the NRA, exportable products are more strongly supported and land endowment, proposed as a proxy for comparative advantage, has a negative effect on market distortions. The use of land-endowment as variable for comparative advantage has severe disadvantages, including missing or falsely reported values, but has also been challenged based on more fundamental and theoretical reasoning. The latter is underpinned by skepticism whether arable land without considering water resources or capital can be seen as proxy for productivity since land on its own is not productive (OECD 2013).

Gawande and Hoekman (2010) look at land endowment, rural population size and political institutions indicators and their effects on agricultural distortions. The results obtained from fixed effects estimation suggest that countries with larger land and labor endowments tend to increase taxation of agricultural production while, on the other hand, more intense electoral competition tends to increase subsidization of the agricultural sector. Additionally to the already mentioned critique put forward against the use of a land variable, another limitation is that the included political institution indicator only allows to distinguish whether there are multiple parties or candidates to choose from at all. Furthermore, the applied empirical measure for political institutions assigns a full democracy status to all countries where the leading party controls just under 75 percent of the seats in the legislature (Cruz, Keefer, and Scartascini 2016). This leads to a loss of information and complicates the interpretation due to missing variation and distinction. Including a larger set of institutional indicators, Olper and Raimondi (2010) claim that the level of agricultural protection increases with transition into stable proportional democracies.

The above mentioned results are based on either the NRA or the RRA as dependent variable to be explained. The OECD database is used for an quantitative analysis in the Agricultural Policy Monitoring and Evaluation Report from 2013 (OECD 2013). There, income, the importance of agriculture in the respective economy and again land endowment are tested as explanatory variables and the findings confirm the existence of policy patterns. Interestingly, the so far discussed literature exclusively rely on static econometric model specifications which only allow to assess the long-run impacts of the respective (policy) variables on market distortions in agricultural markets. Swinnen et al. (2012), by contrast, propose the application of a dynamic model and apply a GMM-based estimator for the OECD's producer support indicator as dependent variable to explain the WTO's effect on agricultural trade liberalization while controlling for income and trade effects. Their findings suggest that the WTO rather supported a shift towards less distorting measures while they cannot find a statistical significant WTO effect on the level of overall agricultural support. Furthermore, they also highlight the persistence in market distortions over time which suggests that the static models applied previously are likley misspecified.

This paper builds on the methodological approach suggested by Swinnen et al. (2012) and formulates a dynamic specification which is estimated by means of different dynamic panel data methods. For the inclusion of control variables we follow the above discussed literature and augment it with information on bilateral trade agreements in order to assess the short- and long-run effects of multilateral versus bilateral trade policies for distortions in agricultural markets. Furthermore, the paper makes use of the larger World Bank database and (mainly) relies on the NRA as the dependent variable of interest.

3 Data, Empirical Specification and Estimation

The data used for the empirical analysis are provided by various different sources including international organizations and databases provided by scientists. The key dependent variable is the indicator for agricultural Ddistortions from the World Bank Estimates of Distortions to Agricultural Incentives Database constructed by Anderson and Nelgen (2013). The nominal rate of assistance (NRA) captures the absolute level of distortions to agricultural prices. *"In other words it measures distortions imposed by governments that create a gap between current domestic prices and the prices that would exist under free markets"* (Anderson et al. 2013). Formally, the NRA is defined as:

$$NRA_{i,t} = \frac{Pd_{i,t} - Pf_{i,t}}{Pf_{i,t}}, \quad (1)$$

where $Pd_{i,t}$ denotes the observed domestic price in country i and year t, $Pf_{i,t}$ is the estimated price in absence of any market distortions. Therefore, the NRA would amount to 0 under a complete free trade environment. It is positive when prices are artificially inflated and negative when producers are net taxed.[4]

The main explanatory variables of interest are the dummy variable URAA and the number of free trade agreements (FTA). The URAA dummy variable takes on the value of one from 1994 onwards for all countries which are among the initial signature states of the Uruguay Agreement and for all other countries starting with the year of their WTO accession. Otherwise the value of the WTO variable is equal to zero. Following Persson (2005) and Swinnen et al. (2012), we additionally control for potential anticipation effects of the Uruguay agreement by including a dummy variable which amounts to one in the six years preceding the official creation of the WTO in 1995 and zero otherwise. For countries joining the WTO in later years (such as e.g., China), the dummy variable amounts to one in the six years prior to their accession and to zero in all other years. FTA expresses the number of (full) bilateral free trade agreements signed by each individual country according to the data collected in DESTA (Dür et al. 2014).[5]

In order to estimate the – ceteris paribus – effects of both multilateral and bilateral trade agreements, the model controls for the effects of several other country characteristics discussed in the previous section and drawing from the recent (empirical) literature. Regarding the other explanatory variables, data for (log) real GDP per capita (i.e., income) measured and agriculture value added per worker are taken from the World Bank World Development Indicators (WDI) database[6]. Agricultural exports and imports are retrieved from the FAOSTAT trade database[7] and are used to construct a relative measure for the agricultural trade balance by dividing net exports in agricultural products by a country's total trade volume in this industry. Furthermore, and in order to control for the relative-income pattern suggested in the literature the specification also accounts for agricultural (log) value added per worker (AVA) as an indicator for comparative advantage offered by the World Bank data. Referring to the findings from previous literature, we also include a measure for political institutions based on political regime characteristics ranging from -10 (strongly autocratic) to +10 (strongly democratic) collected in the Integrated Network for Societal Conflict Research (INSCR) Database (Marshall, Jaggers, and Gurr 2014).[8]

Formally, the resulting dynamic panel data specification to be estimated reads as follows:

[4] The World Bank database also offers the relative rate of aAssistance which additional takes into account market distortions for manufacturing goods and, thus, provides an estimate for either anti- or pro-agriculture biases in a country's support policies. Since this paper focuses on agricultural distortion only, the RRA will not be applied as it is also affected by changes in non-agricultural policy initiatives.

[5] DESTA distinguishes trade agreements by depths of integration with the following categories: (1) partial scope agreement, (2) (full) free trade agreement, (3) customs union, (4) services agreement or (5) framework agreement; no specific provisions.

[6] Available online at http://data.worldbank.org/data-catalog/world-development-indicators.

[7] Available online at http://www.fao.org/faostat/en/home.

[8] Alternatively, on could also control for political orientation based on a government's attitudes towards a market based economy since such economies are likely to signs more bilateral trade agreements and are less reluctant to reduce distorting agricultural support measures. However, due to lack of data availability for the respective measure an inclusion such an indicator would led to the loss of almost half of the available observations and, therefore, we apply the mentioned polity 2 index instead.

$$NRA_{i,t} = \alpha NRA_{i,t-1} + \beta_1 URAA_{i,t} + \beta_2 FTA_{i,t} + \mathbf{x}_{i,t}\gamma + \mu_i + \zeta_t + \epsilon_{i,t} \quad (2)$$

where $NRA_{i,t}$ is the nominal rate of assistance in country i in time t, μ_i denotes a country-specific effect which captures unobserved but time-invariant characteristics and ζ_t are common time-specific fixed effects (such as e.g., time trends in world food prices). $\epsilon_{i,t}$ is the reminder error term which is assumed to be independently but not necessarily identically distributed (i.e., we allow for heteroscedasticity). $\mathbf{x}_{i,t} = \left[ln(INC_{i,t}), ln(INC_{i,t})^2, ATB_{i,t}, ln(AVA_{i,t}), DEM_{i,t}, URAApt_{i,t}\right]$ constituting a vector of all country-time-specific control variables with the corresponding vector of estimable parameters γ. The inclusion of the squared income term aims at capturing likely non-linear effects in the relationship between GDP per capital and agricultural market distortions identified in the previous literature and the Figures discussed in Section 2.1.

Furthermore, the parameter α captures the persistence in agricultural distortions over time and will be used to calculate the long-run effects of the covariates of main interest. As discussed above, these are the $URAA_{i,t}$ indicator and the number of country-time-specific (full) bilateral free trade agreements in force ($FTA_{i,t}$). Their respective estimable parameter values are denoted by β_1 and β_2, respectively, which measure the short-run (direct) impacts of multilateral and bilateral trade policies on agricultural market distortions. The long-run effect of an additionally signed trade agreement is, for example, given by $\frac{\beta_2}{1-\alpha}$.

Equation (2) is estimated by means of various different panel data estimators. They commonly eliminate μ_i by either applying the within-transformation or first-differencing the model specification. The different estimators proposed vary in their assumptions relevant for identifying unbiased estimators for the parameters associated with the lagged dependent variable (and all other predetermined characteristics) and the number of bilateral trade agreements. In a dynamic stochastic equation, the immediate problem that $NRA_{i,t-1}$ is correlated with the error term when applying simple ordinary least squares (OLS) to within-transformed data gives rise to the dynamic panel bias as pointed out first by Nickell (1981). Judson and Owen (1999) discuss this issues and compares the performance of different dynamic panel data estimators including e.g., simple and bias-correcting OLS-estimators and generalized methods of moments (GMM) estimators using a Monte Monte Carlo simulation exercise. Based on their findings, the paper concludes that when the time-dimension T gets relatively large (as compared to the number of cross-section units N), a GMM-estimation may not be practical. Instead Judson and Owen (1999) find that for balanced panal datasets the bias-correcting approach proposed by Bun and Kiviet (2003) outperforms other estimators. An extension of this approach for unbalanced panel data (as are used in this exercise) is offered by Bruno (2005).

Furthermore, one could raise the objection of an apparent endogeneity problem concerning the number of free trade agreements a country agrees on, as unobserved and time-variant preferences regarding the liberalization of markets might be correlated with both bilateral trade policy activities and the distortion in agricultural markets. This issue will be addressed by applying two instruments for the number of bilateral free trade agreements a country agreed on: The first one is based on empirical research on the determinants of the likelihood to sign free trade agreements which suggests that trade agreements are highly contagious and induce domino effects (Baier, Bergstrand, and Mariutto 2014; Baldwin and Jaimovich 2010; Egger and Larch 2008). Accordingly, we argue that bilateral trade policy activities of neighboring economies exogenously pressure the respective countries to also engage in trade agreement negotiations and, thus, we instrument the number of free trade agreement signed by each country itself by the average number of trade agreements signed by a country's (bordering) neighbors excluding all bilateral agreements from the neighboring countries with country i. Based on a similar reasoning and following the approach suggested by Cadot et al. (2010), the second instrument applied is the number of military alliances the country has joined over time. In a similar vein as for the first instrument, integration into various military alliances is expected to foster economic integration as well, making the signing of bilateral free trade agreements more likely. The respective data for memberships in military alliances is taken from Gibler (2008).

Table 1: Descriptive statistics

	Obs	Mean	Std. Dev.	Min	Max
NRA	1,897	0.223	0.566	-0.852	4.321
$URAA$	1,897	0.588	0.492	0	1
FTA	1,897	21.255	22.308	0	93
INC	1,897	12,755.65	15,363.20	123.730	69,094.75
AVA	1,897	11,056.03	15,181.39	122.748	143,036.50
ATB	1,897	0.017	0.430	-0.929	0.937
$URAApt$	1,897	0.767	0.423	0	1
DEM	1,897	5.098	6.135	-9	10

Table 1 provides am overview on the data sample utilized for the empirical analysis by reporting simple descriptive statistics. The full sample contains 1,897 observations based on 76 different countries and capturing the years ranging from 1980 to 2011.[9] The dependent variable of interest, NRA takes on a sample mean of approximately 0.23 implying that, on average, realized prices in agricultural markets are around 23% above their non-distortionary values. For about one-third of all observations, we observe negative realizations of the NRA indicating a relative overtaxation of agricultural products in these cases. In the remaining two-thirds of all observations agricultural products are, however, net subsidized translating into above free market prices.

With regard to the main trade policy variables of interest, in around 59% of all observations, the countries are member states of WTO and thus took part in the multilateral trade liberalization policies agreed on in the Uruguay round. In terms of bilateral trade relationships we observe a relatively large variation in signed agreements across countries and time (see also Figure 3). On average, each country agreed on approximately 21 bilateral trade agreements over the whole sample period. Germany most actively signs bilateral trade agreements and at the end of our sample period had 93 such agreements in force. In 2011 all countries available in the sample had agreed on at least one bilateral trade agreement while in 2010 Ecuador and Nigeria formed the last two countries without having any bilateral trade agreement (as defined above) in force.

4 Estimation Results and Discussion

Table 2 reports the main results from estimating Equation (2) using various different (dynamic) panel data estimators. Column 1 displays the parameters obtained from simple within-transformed (fixed effects) estimation. The results in Columns 2 and 3 explicitly account for the Nickel bias by applying the bias-corrected least squares dummy (LSDV) estimator (Column 2) and a GMM-based approach

[9] A detailed description of the sample composition is offered in Table A1 in the appendix.

(Column 3). In column (4) we turn to the potential endogeneity of the number of bilateral trade agreements signed and thus provide results from fixed-effects instrumental variables (IV) estimation, applying the average number of bilateral trade agreements signed by a country's neighboring economies excluding the direct bilateral trade agreements signed as our instrument. Due to the fact that multi-coliniarity is a more severe issue in instrumental variables estimation, we rely on a more parsimonious specification when applying the GMM estimator excluding the squared-term of (log) income.[10] As internal instruments and based on the Hanson J-test we use the levels of the third and fourth lag thereby restricting the number of instruments to remain relatively small. As additional instruments, we use both the average number of free trade agreements signed by neighboring economies and the number of military alliances in which the country participates. We further assume (log) income, the democratization score and WTO membership to be predetermined.

Table 2: Main estimation results

	Fixed-Effects (1)	LSDV-corrected (2)	GMM (3)	Fixed Effects IV (4)
$NRA_{i,t-1}$	0.719^{***}	0.779^{***}	0.834^{***}	0.712^{***}
	(0.016)	(0.017)	(0.058)	(0.045)
$URAA_{i,t}$	-0.023	-0.021	-0.011	-0.022
	(0.024)	(0.038)	(0.121)	(0.025)
$FTA_{i,t}$	-0.003^{***}	-0.002^{***}	-0.002^{*}	-0.004^{***}
	(0.001)	(0.001)	(0.001)	(0.001)
$ln(INC_{i,t})$	0.429^{***}	0.293^{***}	0.213^{**}	0.378^{***}
	(0.116)	(0.091)	(0.091)	(0.120)
$ln(INC_{i,t})^2$	-0.025^{***}	-0.017^{***}	-	-0.022^{***}
	(0.008)	(0.006)	-	(0.008)
$ln(AVA_{i,t})$	-0.038	-0.028	-0.209^{**}	-0.040
	(0.039)	(0.023)	(0.094)	(0.027)
$ATB_{i,t}$	-0.090^{**}	-0.083^{***}	-0.124	-0.084^{***}
	(0.041)	(0.030)	(0.081)	(0.028)
$URAApt_{i,t}$	-0.033	-0.032	-0.042	-0.033
	(0.029)	(0.044)	(0.200)	(0.032)
$DEM_{i,t}$	0.005^{***}	0.004^{***}	0.003	0.005^{***}
	(0.002)	(0.001)	(0.003)	(0.001)
No. of instruments	-	-	89	1
Partial F-statistic	-	-	-	1010.48
Hanson J-test: p-value	-	-	0.778	-
AR(1): p-value	-	-	0.000	-
AR(2): p-value	-	-	0.144	-
R_a^2	0.688	-	-	0.675
Obersavtions	1,897	1,897	1,765	1,895

Notes: Clustered standard errors at the country-level in parentheses. *, **, *** . . . significant at 10%-, 5%- and 1%-level, respectively. In column (4) $FTA_{i,t}$ is instrumented with the average number of bilateral trade agreements signed by the neighboring economies excluding the direct bilateral trade relationships.

Starting the discussion with the results regarding the control variables first, the different estimators provide remarkably similar results. A country's average income level measured in terms of GDP per capita exhibits a robust positive but diminishing impact on the NRA indicated by the significant positive parameters associated with linear income term and the significant negative parameters associated

[10] When applying the GMM-approach to the full-specifiction all control variables lose their statistical significance, while the estimates for the lagged NRA and for FTA are hardly affected. These results are available from the authors upon request.

with squared income (see Columns 1, 2 and 4). Equipped with the estimated parameters associated with both different income terms and the distribution of (log) income within the sample (see Table 1), we can calculate the overall income effects for different income levels. Accordingly, the overall effect of income on agricultural market distortions remains positive throughout the whole income distribution and across the different estimators applied. Market distortions in agricultural markets are thus increasing with a country's income but at a decreasing rate.

Furthermore and also in line with the previous literature, Table 2 reveals significant negative effects of a country's agricultural trade balance implying that countries which are net exports of agricultural products are less engaged in providing political support in order to distort agricultural markets. This effect again remains very stable across most of the different estimation methods applied. In a similar vein, more democratic political regimes tend to systematically distort agricultural markets. This is indicated by the positive and statistically highly significant parameter estimates which are, however, not very sizeable in economic terms (see Columns 1, 2 and 4). Finally, variation in labor productivity in the agricultural industry, as measured by (log) value added per employee is not able to additionally explain the remaining variation in agricultural market distortions. The only exception for this is offered by the results from the GMM-estimator for which we identify a statistically significant negative parameter estimate associated with with $ln(AVA_{i,t})$.

Turning our attention to the persistence in agricultural market distortions as governed by the parameter estimates corresponding to lagged $NRA_{i,t-1}$, we are not able to identify a substantial Nickel Bias in Columns (2) and (4). The parameter values obtained from simple within OLS and fixed-effects IV are only slightly smaller as the ones obtained from the bias-corrected OLS estimator and from system GMM. Due to the relatively large time-dimension (capturing 33 years) together with a relatively small number of cross-sectional units, these findings are consistent with the simulation exercise provided in Judson and Owen (1999). For the subsequent discussion of the findings associated with the variables of main interest this implies that we can safely rely on the simpler estimators proposed. Furthermore, these results do also not vary substantially across the different estimators which gives support to the robustness of our main findings.

Focusing on the short-run effects of multilateral trade negotiations first, our results point to a limited contribution of the WTO membership for decreasing market distortions in international agricultural markets. Neither direct membership captured by $URRAA_{i,t}$ nor potential membership anticipation (measured via $URAApt_{i,t}$) are indicted to directly affect the observed $NRA_{i,t}$. The estimated parameters are all insignificant throughout. This finding is well in line with the evidence provided by Swinnen et al. (2012). In contrast to this we are able to identify statistically significant and negative effects for bilateral trade agreements. Accordingly, an increase in the number of bilateral trade agreements a country has in force reduces market distortions in its agricultural market. In quantitative terms and at a first glance the estimated effect seems rather small varying between -0.002 (obtained from the bias corrected OLS estimator and the GMM estimator) and -0.004 when accounting for potential endogeneity of the number of bilateral trade agreements in the fixed-effects IV estimator. However, when accounting for the variation in the number of bilateral trade agreements signed, the effect of bilateral trade agreements turn out to be quite sizeable. Accordingly and in the short-run only, a one standard deviation increase in the number of signed trade agreements (i.e., 22.3) decreases the NRA by a magnitude of between -0.045 and -0.089. In the long-run the estimated effects of a one standard deviation increase in the number of bilateral trade agreements are substantially larger and amount to -0.202 and -0.309. Put differently, in the long-run one standard deviation in the number of bilateral trade agreements is able to explain approximately one-half of the standard deviation in the observed NRA distribution (which amounts to 0.566, see Table 1). Based on our estimates, bilateral trade policy is identified as a crucial additional determinant of the level of market distortions in international agricultural markets which has been mainly ignored in the literature so far. While multilaterilism seems to be ineffective in reducing political support for agricultural products both in the short- and long-run, bilateral trade agreements seem to systematically contribute to the removal of such barriers for agricultural market liberalization.

Due to likely endogeneity of the number of bilateral trade agreements (and the absence of strong evidence for a Nickel bias) we prefer the fixed-effects IV estimator with its corresponding results being displayed in Column (4) and subsequently further investigate the robustness of our main finding relying on this estimator only. The respective estimation results are reported in Columns (1) to (8) of Table 3. Column (1) uses the lagged value of the average number of neighboring FTAs, Column (2) uses both the lagged and the contemporaneous realizations of this instrument. Columns (3) and (4) rely on four years lagged values of the average free trade agreements signed by the neighboring economies excluding as usual all direct bilateral relationships. Column (4) adds the fours years lag of the own number of signed bilateral trade agreements as an additional instrument. In column (5) we provide a reduced form estimate for $FTA_{i,t}$ by using WTO membership as an additional instrument. This choice is motivated by insignificant direct effects identified for WTO membership in the baseline estimates provided in Table 2. Accordingly, one could argue that taking part in the WTO might shift the probability of additionally signing bilateral free-trade agreements which would qualify $URAA_{i,t}$ to be a useful instrument. In Columns (6) and (7) we alternatively apply the number of military alliances in which the countries participate as instrument. In the former, we only rely on this instrument while Column (7) provides the results based on the joint inclusion of both instruments discussed in Section 3. The final column re-estimates our baseline specification from Column (4) of Table 2 for the smaller sample based on the OECD PSE data sample and relying on the NAC instead of the NRA as the outcome variable of interest.

The partial F-statistics reported at the bottom of Table 3 commonly point to the relevance of the instruments applied (similar to the one reported in Column 4 of Table 2). These values are all well above the rule of thumb criteria requiring the partial F to be above 10 (Stock, Wright, and Yogo 2002). Furthermore, for all cases in which we include more than just one instrument, the p-values associated with Hanson J-tests support our assumption concerning the exogeneity of the used instruments. Across all different specification applied, the effect of our main variable of interest remains remarkably stable. Accordingly and in line with our baseline results, signing one additional bilateral trade agreement reduces the NRA of a participating economy by 0.004 to 0.005. This effect only loses its statistically significance in Column (8) where we alternatively rely on the OECD PSE data instead of the agricultural distortion database. However, given the small number of less than 500 observations and the small country coverage provided in the PSE database, this result does not come as a big surprise. In qualitative terms the FTA effect from Column (8) is very similar to the ones identified in the specifications lending additional support to our main findings. Taking all the findings from the alternative estimators and specifications together, the empirical analysis carried out in this section suggests a statistically significant and economically relevant impact of bilateral trade agreements for reducing market distortions in agricultural markets.

Table 3: Robustness checks: Estimation results

	FEIV 1 (1)	FEIV 2 (2)	FEIV 3 (3)	FEIV 4 (4)	FEIV 5 (5)	FEIV 6 (6)	FEIV 7 (7)	FEIV 8 (8)
$NRA_{i,t-1}$	0.710***	0.712***	0.697***	0.701***	0.711***	0.676***	0.683***	-
	(0.045)	(0.045)	(0.049)	(0.048)	(0.045)	(0.019)	(0.037)	-
$NAC_{i,t-1}$	-	-	-	-	-	-	-	0.611***
	-	-	-	-	-	-	-	(0.069)
$URAA_{i,t}$	−0.024	−0.024	−0.026	−0.027	-	−0.020	−0.021	−0.038
	(0.026)	(0.026)	(0.027)	(0.027)	-	(0.026)	(0.025)	(0.033)
$FTA_{i,t}$	−0.004***	−0.004***	−0.004***	−0.004***	−0.004***	−0.005*	−0.004***	−0.005
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.003)	(0.001)	(0.004)
$ln(INC)_{i,t}$	0.364***	0.374***	0.405***	0.427***	0.375***	0.426***	0.460***	1.331***
	(0.121)	(0.120)	(0.143)	(0.141)	(0.119)	(0.132)	(0.111)	(0.463)
$ln(INC)^2_{i,t}$	−0.021***	−0.022***	−0.023**	−0.024***	−0.022***	−0.025***	−0.027***	−0.095***
	(0.008)	(0.008)	(0.009)	(0.009)	(0.008)	(0.009)	(0.008)	(0.032)
$ln(AVA)_{i,t}$	−0.040	−0.040	−0.043	−0.043	−0.041	−0.043*	−0.043*	−0.041
	(0.027)	(0.027)	(0.031)	(0.031)	(0.027)	(0.024)	(0.024)	(0.093)
$ATB_{i,t}$	−0.080***	−0.082***	−0.080**	−0.083**	−0.080***	−0.084***	−0.086***	−0.340***
	(0.028)	(0.028)	(0.033)	(0.033)	(0.027)	(0.026)	(0.025)	(0.092)
$URAApt_{i,t}$	−0.027	−0.027	−0.039	−0.039	-	−0.028	−0.027	0.039
	(0.033)	(0.033)	(0.034)	(0.034)	-	(0.031)	(0.031)	(0.038)
$DEM_{i,t}$	0.005***	0.005***	0.005***	0.005***	0.005***	0.004***	0.004***	0.008
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.005)
No. of instruments	1	2	1	2	2	1	2	1
Partial $F-$statistic	788.11	502.94	437.50	325.65	514.57	41.51	449.08	47.83
Hanson J-test: p-value	-	0.3310	-	0.2502	0.3679	-	0.7296	-
R^2_a	0.674	0.674	0.675	0.676	0.675	0.664	0.666	0.637
Obersavtions	1,893	1,893	1,747	1,747	1,893	1,761	1,761	499

Notes: Clustered standard errors at the country-level in parentheses. *, **, *** ...significant at 10%-, 5%- and 1%-level, respectively. The observations of Russia and Ukraine were dropped prior to the year of 1993 to avoid disturbances by the break-up of the Soviet Union. In columns 1 to 4 $FTA_{i,t}$ is instrumented by (1) one year lagged average free trade agreements of the neighbors, (2) contemporaneous and one year lagged neighboring average agreements, (3) 4-years lagged average number of bilateral trade agreements signed by the neighboring economies, (4) 4-year lagged own and neighboring average agreements. Column (5) uses $URAA$ as an additional instrument instead of an explanatory variable, (6) uses the military alliances participation as the only instrument, (7) uses military alliances and the contemporaneous number of average neighboring FTAs as instruments. Column (8) re-estimates the baseline from Column (4) in Table 2 with the OECD PSE data sample.

5 Conclusions

This paper builds on the empirical literature on distortions in agricultural markets and puts a specific focus on the role of international trade negotiations for shaping the institutional framework for agricultural policies. In particular, we study the short- and long-run effects of multilateral and bilateral trade agreements for reducing market distortions in agricultural markets. For separating the short-run from the long-run effects of trade policy, we formulate a dynamic panel data equation which is theoretically motivated by political economy reasoning for agricultural policy. The resulting econometric specification is then estimated by means of various different panel data methods including simple fixed-effects OLS and IV estimators, the least square dummy variable corrected (LSDVC) estimator and a more general GMM-based approach. The empirical estimates based on the alternative econometric methods applied point to a remarkably robustness of the main results and further do not indicate any substantive dynamic panel data bias. Given the relatively long time-dimension of 32 years in the dataset the latter result does not come as a big surprise and is well in line with econometric research on the performance of different dynamic panel data estimators.

With regard to previous empirical literature on market distortions in agricultural markets, we are able to confirm the main findings using an enlarged and updated dataset capturing 76 economies for the years ranging from 1980 to 2011. In particular, a growing per capita income in a country leads to an increase in political support for agricultural producers inducing positive price distortions as measured by the World Bank's nominal rate of assistance (NRA). Our estimates indicate a non-linear relationship between income and market distortions in agricultural markets, indicating that price distorting subsidies are getting smaller with income. By contrast, low income countries generally still tend to depress producer prices by overtaxing agricultural products. (see, e.g., Table A1 in the Appendix). High income countries tend to redesign their political support measures such that they are less price distorting which is also consistent with the non-linear relationship identified empirically. Furthermore, economies with a positive net trade balance tend to grant lower support levels or even depress prices for exporters and producers while – ceteris paribus – countries with more sophisticated democratic institutions tend to provide more price distorting support measures for agricultural products.

Putting the focus on the impact of multilateral and bilateral trade negotiations on agricultural market distortions, we obtain heterogeneous and potentially policy relevant results. First of all, we are not able to identify any significant liberalization effects stemming from the WTO's Uruguay agreement for the sample country at hand. The estimates tend to be negative throughout, which would be associated with decreasing market distortions, but lack statistical significance in all specifications and across the various alternative econometric estimators applied. This finding is well in line with the results provided by Swinnen et al. (2012) who do not find significant distortion reducing WTO effects using the OECD's Producer Support Estimates data instead. Putting this finding into perspective, economists tend to agree that the GATT and the WTO have been successful in liberalizing trade for manufacturing goods. However, the verdict is less conclusive when it comes to the agricultural sector and in particular skepticism has been articulated with regard to the first major trade agreement concentrating on agricultural markets - the Uruguay Round Agreement on Agriculture (Swinnen et al. 2012). The OECD (2001) called the Uruguay Round Agreement on Agriculture a watershed, in the sense that agriculture was finally subjected to multilateral rules and disciplines, but only finds moderate trade barrier reduction associated with this agreement. Furthermore, this agreement can be seen as an international framework for standardizing the evaluation of tariffs and defining common areas of focus for specific reforms (Cahill and Brooks 2001; OECD 2001), but given all the available evidence has yet not systematically contributed to the reduction of agricultural market distortions across the world.

Turning the attention to the key finding of this paper, the phenomenon triggered in the 1990s of a surging number of signed bilateral trade agreements is without a doubt one of the most significant changes in international trade relationships over this period of time and not surprisingly influences market access and the level of agricultural price distortions. In our exercise the variation in the signed bilateral trade agreements by countries systematically explains differences in agricultural market distortions. In our preferred specification (Column 4 of Table 2), a one standard deviation increase in the number of bilateral agreements in force (i.e., 22.3 agreements) decreases the nominal assistance coefficient by -0.089 and -0.309 in the short- and long-run, respectively. These effects are also economically relevant indicating that bilateral trade agreements substantially contribute to fostering efficiency in agricultural markets. Putting all empirical findings from this paper together, regional trade policy and bilateralism seem to constitute be the most effective trade policy tools for reducing market distortions in agricultural markets.

From the signing of the North American Free Trade Agreement NAFTA, the various bilateral EU agreements with other countries to the preliminary entry into force of the Trans-Pacific Partnership (TPP) and the long lasting attempts to conclude on the Transatlantic Trade and Investment Partnership (TTIP) between the USA and the EU, regionalism has been seen as the prime vehicle for further fostering international market integration by reducing remaining barriers to trade (Matthews 2013). With Great Britain's decision to leave the European Union and the election of Donald Trump as the 45th president of the United States both taking place in 2016, the political views on regionalism and bilateral trade agreements seem to have been substantially reversed and a lot of skepticism towards free trade and globalization has been articulated. Taking our estimation results literally, policies which would roll back bilateral trade agreements are expected to increase market distortions in agricultural markets inducing price rises for the consumers of such goods. From a distributional point of view this would likely hurt the lowest income groups the most, as relying on Engel's law these households use a larger share of total expenditures for the consumption of agricultural products.

References

Adams, Leonard P. (2013). *Agricultural Depression and Farm Relief in England 1813-1852*. Abingdon: Routledge.

Anania, Giovanni, Mary E. Bohman, Colin Carter, and Alex F. McCalla (2004). *Agricultural Policy Reform and the WTO. Where are we heading?* Cheltenham: Edward Elgar Publishing.

Anderson, Kym (2009). Distorted agricultural incentives and economic development: Asia's experience. *The World Economy* 32(3), 351–384.

Anderson, Kym (2010). *The Political Economy of Agricultural Price Distortions*. New York: Cambridge University Press.

Anderson, Kym and Signe Nelgen (2013). *Research - Estimates of distortions to agricultural incentives, 1955-2011 (updated June 2013)*. www.worldbank.org/agdistortions. Online; accessed 15 May 2016.

Anderson, Kym, Gordon Rausser, and Johan F.M. Swinnen (2013). Political economy of public policies: insights from distortions to agricultural and food markets. *Journal of Economic Literature* 51(2), 423–477.

Bagwell, Kyle, Chad P. Bown, and Robert W. Staiger (2016). Is the WTO passé? *Journal of Economic Literature* 54(4), 1125–1231.

Baier, Scott L., Jeffrey H. Bergstrand, and Ronald Mariutto (2014). Economic determinants of free trade agreements revisited: Distinguishing sources of interdependence. *Review of International Economics* 22(1), 31–58.

Baldwin, Richard and Dany Jaimovich (2010). Are free trade agreements contagious? *NBER Working Paper* No. 16084.

Brooks, Jonathan (1996). Agricultural policies in OECD countries: What can we learn from political economy models? *Journal of Agricultural Economics* 47(1-4), 366–389.

Bruno, Giovanni S.F. (2005). Estimation and inference in dynamic unbalanced panel data models with a small number of individuals. *CESPRI Working Paper No. 165.*

Buchanan, James M. and Gordon Tullock (1962). *The Calculus of Consent.* Vol. 3. Ann Arbor: University of Michigan Press.

Bun, Maurice and Jan Kiviet (2003). On the diminishing returns of higher-order terms in asymptotic expansions of bias. *Economics Letters* 79(2), 145–152.

Cadot, Olivier, Marcelo Olarreaga, and Jeanne Tschopp (2010). Trade agreements and trade barrier volatility. In: Anderson, Kym. *The Political Economy of Agricultural Price Distortions.* New York: Cambridge University Press, 332–357.

Cahill, Carmel and Jonathan Brooks (2001). Why agricultural trade liberalisation matters. *OECD Observer* 229, 49–50.

Canada, Government of (2016). *North american free trade agreement (NAFTA) - Agriculture and Agri-Food Canada (AAFC).* http://www.agr.gc.ca/eng/?id=1384265676755. Online; accessed 11 November 2016.

Cruz, Cesi, Philip Keefer, and Carlos Scartascini (2016). *The database of political institutions 2015 (DPI2015).* https://publications.iadb.org/handle/11319/7408. Online; accessed 10 November 2016.

De Gorter, Harry and Yacov Tsur (1991). Explaining price policy bias in agriculture: The calculus of support-maximizing politicians. *American Journal of Agricultural Economics* 73(4), 1244–1254.

Downs, Anthony (1957). *An Economic Theory of Democracy.* New York: Harper and Row.

Dür, Andreas, Leonardo Baccini, and Manfred Elsig (2014). The design of international trade agreements: Introducing a new dataset. *The Review of International Organizations* 9(3), 353–375.

Egger, Peter and Mario Larch (2008). Interdependent preferential trade agreement memberships: An empirical analysis. *Journal of International Economics* 76(2), 384–399.

Engel, Ernst (1857). Die Produktions- und Konsumptionsverhältnisse des Königreichs Sachsen. *Zeitschrift des Statistischen Bureaus des Königlich Sächsischen Ministeriums des Innern* 8, 1–54.

Ethier, Wilfred J. (1998). Regionalism in a multilateral world. *Journal of Political Economy* 106(6), 1214–1245.

Freund, Caroline (2000). Multilateralism and the endogenous formation of preferential trade agreements. *Journal of International Economics* 52(2), 359–376.

Gawande, Kishore and Bernard Hoekman (2010). Why governments tax or subsidize trade: Evidence from agriculture. In: Anderson, Kym. *The Political Economy of Agricultural Price Distortions.* New York: Cambridge University Press, 241–277.

Gibler, Douglas M. (2008). *International Military Alliances, 1648-2008.* Washington D.C.: CQ Press.

Judson, Ruth A. and Ann Owen (1999). Estimating dynamic panel data models: A guide for macroeconomists. *Economics Letters* 65(1), 9–15.

Marshall, Monty G., Keith Jaggers, and Ted R. Gurr (2014). *Polity IV project: Political regime characteristics and transitions, 1800–2013. INSCR.* http://www.systemicpeace.org/inscrdata.html. Online; accessed 29 November 2016.

Masters, William A. and Andres F. Garcia (2010). The political economy of agricultural price distortion and stabilization. In: Anderson, Kym. *The Political Economy of Agricultural Price Distortions*. New York: Cambridge University Press, 215–240.

Matthews, Alan (2013). Doha negotiations on agriculture and future of the WTO multilateral trade system. *Mimeo*.

Nickell, Stephen (1981). Biases in dynamic models with fixed effects. *Econometrica*, 1417–1426.

OECD (2001). *The Uruguay Round Agreement on Agriculture*. Paris: OECD Publishing.

OECD (2013). *Agricultural Policy Monitoring and Evaluation 2013*. Paris: OECD Publishing.

OECD (2016). *Agricultural Policy Monitoring and Evaluation 2016*. Paris: OECD Publishing.

Olper, Alessandro and Valentina Raimondi (2010). Constitutional rules and agricultural policy outcomes. In: Anderson, Kym. *The Political Economy of Agricultural Price Distortions*. New York: Cambridge University Press, 358–392.

Olson, Mancur (1965). *The Logic of Collective Action*. Vol. 124. Cambridge: Harvard University Press 2009.

Ornelas, Emanuel (2005). Trade creating free trade areas and the undermining of multilateralism. *European Economic Review* 49(7), 1717–1735.

Persson, Torsten (2005). Forms of democracy, policy and economic development. *NBER Working Paper* No. W11171.

Schonhardt-Bailey, Cheryl (2006). *From the Corn Laws to Free Trade: Interests, Ideas, and Institutions in Historical Perspective*. Cambridge: MIT Press.

Stock, James H., Jonathan H. Wright, and Motohiro Yogo (2002). A survey of weak instruments and weak identification in generalized method of moments. *Journal of Business & Economic Statistics* 20(4), 518–529.

Swinnen, Johan F.M. (2010). Political economy of agricultural distortions: The literature to date. In: Anderson, Kym. *The Political Economy of Agricultural Price Distortions*. New York: Cambridge University Press, 81–104.

Swinnen, Johan F.M., Alessandro Olper, and Thijs Vandemoortele (2012). Impact of the WTO on Agricultural and Food Policies. *The World Economy* 35(9), 1089–1101.

UK Parliament (2016). *About Parliament - Living heritage: Free trade and protection. Contains Parliamentary information licensed under the Open Parliament Licence v3.0.* www.parliament.uk/about/living-heritage/transformingsociety/tradeindustry/importexport/overview/freetrade. Online; accessed 10 June 2016.

Appendix: Country Coverage and Sample Composition

Table A1: Country coverage and sample overview

Country-code	**First obs.**	**Last obs.**	**NRA initial year**	**NRA latest year**
ARG	1980	2010	−0.059	−0.282
AUS	1980	2009	0.036	0.000
AUT	1980	2011	0.185	0.019
BEL	1980	2011	0.547	0.040
BEN	1980	2005	−0.006	0.003
BFA	1980	2010	−0.032	−0.237
BGD	1980	2009	−0.099	−0.425
BGR	1992	2011	−0.318	0.010
BRA	1980	2010	−0.318	0.009
CAN	1980	2011	0.183	0.097
CHE	1980	2011	2.297	0.538
CHL	1980	2011	0.056	0.002
CHN	1981	2010	−0.533	0.161
CIV	1980	2009	−0.406	−0.331
CMR	1980	2009	−0.156	−0.068
COL	1980	2010	−0.049	0.387
CYP	2005	2011	0.167	0.035
CZE	1992	2011	0.182	0.016
DEU	1980	2011	0.559	0.020
DNK	1980	2011	0.583	0.014
DOM	1980	2010	−0.255	0.430
ECU	1980	2010	−0.071	−0.248
EGY	1980	2010	−0.360	0.028
ESP	1980	2011	−0.184	0.021
EST	1992	2011	−0.390	0.019
ETH	1981	2010	−0.100	−0.224
FIN	1980	2011	0.401	0.023
FRA	1980	2011	0.504	0.018
GBR	1980	2011	0.585	0.032
GHA	1980	2010	−0.231	−0.052
GRC	1981	2011	0.537	0.020
HUN	1992	2011	0.158	0.017
IDN	1980	2010	−0.035	0.293
IND	1980	2010	−0.089	0.139
IRL	1980	2011	0.604	0.029
ISL	1980	2011	2.326	0.532
ISR	1995	2011	0.204	0.134
ITA	1980	2011	0.359	0.024
JPN	1980	2011	1.006	0.997
KAZ	1995	2011	0.177	0.088
KEN	1980	2010	−0.380	−0.106

Continued on next page

Table A1: *Continued from previous page*

Country-code	**First obs.**	**Last obs.**	**NRA initial year**	**NRA latest year**
KOR	1980	2011	1.120	1.046
LKA	1980	2010	−0.334	−0.164
LTU	1992	2011	−0.459	0.023
LVA	1992	2011	−0.460	0.023
MAL	2005	2011	0.145	0.031
MAR	1995	2009	0.595	0.471
MDG	1980	2010	−0.347	−0.150
MEX	1980	2011	0.052	0.029
MLI	1980	2010	−0.050	−0.290
MOZ	1980	2010	−0.450	0.364
MYS	1980	2010	−0.172	0.343
NGA	1980	2010	0.094	−0.006
NIC	1991	2010	−0.081	−0.125
NLD	1980	2011	0.748	0.039
NOR	1980	2011	2.785	0.842
NZL	1980	2011	0.159	0.007
PAK	1980	2010	−0.201	−0.086
PHL	1980	2010	−0.155	0.253
POL	1992	2011	−0.039	0.034
PRT	1980	2011	0.304	0.029
ROM	1992	2011	0.090	0.017
RUS	1992	2010	−0.483	0.152
SDN	1980	2010	−0.360	−0.004
SEN	1980	2005	−0.432	0.160
SVK	1992	2011	0.243	0.014
SVN	1992	2011	0.643	0.034
SWE	1980	2011	0.738	0.021
TCD	1980	2005	−0.053	0.005
TGO	1980	2010	−0.014	−0.134
THA	1980	2010	−0.121	0.104
TUR	1980	2011	−0.237	0.249
TWN	1980	2011	0.286	0.234
TZA	1980	2010	−0.599	0.046
UGA	1980	2011	−0.094	−0.280
UKR	1992	2010	−0.475	0.014
USA	1980	2011	0.078	0.010
VNM	1986	2005	−0.111	0.112
ZAF	1980	2010	0.148	0.006
ZMB	1980	2005	−0.491	0.070
ZWE	1980	2005	−0.522	−0.224