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WTO Membership and the Extensive Margin of World Trade: New Evidence

by

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Abstract

Recent literature has argued that, contrary to the results of a seminal paper by Rose (2004), WTO membership does promote bilateral trade, at least for developed economies and if membership includes non-formal compliance. We review the literature in order to identify open issues. We then develop the simplest possible "corner-solutions" version of the gravity model which serves as a framework to readdress these issues. We focus on the extensive margin of trade that separates positive-trade from zero-trade country pairs. We argue that the model can be consistently estimated using Poisson pseudo-maximum-likelihood methods with exporter and importer fixed effects. We account for coding issues and the potential heterogeneity of the WTO membership which recent contributions have stressed. While we find that WTO membership increases the likelihood that a given country pair trades, we do not find that the extensive margin has a strong and systematic effect on the average trade-creating potential of the WTO.

JEL classification: F12, F13

Keywords: gravity approach, WTO, monopolistic competition, real trade costs

Outline

Non-Technical Summary

According to received wisdom, the GATT/WTO has been a causal factor for an impressive postwar record of trade-induced economic growth and development. If this view is correct, the data should tell us that GATT/WTO membership has had a positive impact on the magnitude of trade. Somewhat surprisingly, a study by Rose (2004) has found no robust empirical evidence of a significant trade-promoting effect of WTO membership, using the gravity model to control for other determinants of bilateral trade. However, subsequent literature has argued that a positive effect does emerge for the group of industrial countries, and if WTO membership is defined so as to include informal compliance with WTO rules, in addition to formal membership.

In this literature, researchers have mostly restricted their analysis to country-pairs with non-zero trade, thus looking at the amount of trade between country pairs where a trading relationship already exists (intensive margin). However, world trade has increased also through newly established trading relationships (extensive margin). Econometric theory should lead us to expect that ignoring evidence on the extensive margin introduces a downward bias in the estimated trade effect of WTO membership. Studies that do include the extensive margin tend to corroborate this expectation.

However, doubts remain. In this paper, we first review the literature in order to identify open methodological issues. One of the problems is that the mainstream gravity model of bilateral trade does not allow for zero bilateral trade. We develop a suitably parsimonious model that is able to explain zero trade and, thus, movements at the extensive margin of world trade. The model stresses country-specific fixed costs that exporting firms have to incur when entering a new market. We explore the implication that this has for the gravity equation of bilateral trade, with particular emphasis on the role of WTO membership.

We then estimate this model, duly taking into account methodological concerns that have been raised in recent literature. These revolve around the appropriate definition of WTO membership, as well as unobserved heterogeneity across countries that may contaminate econometric estimates of the WTO membership effect on trade. Moreover, on grounds of data reliability we caution against the use of time-series evidence on the extensive margin, and we mainly focus on cross-country variance.

Running Probit estimation, we do find some evidence for WTO membership to raise the odds that countries trade with each other at all, but the effect is by no means robust across country groups and time. Nor does a broader definition of membership that includes de facto participation make much of a difference. The asymmetry, both across country groups and types of membership, changes erratically over sub-periods considered.

Running Poisson pseudo-maximum-likelihood estimation gives rise to an even bleaker picture for formal WTO membership than do conventional OLS estimates relying on non-zero trade observations. It is only for the aftermath of the Uruguay-round that we obtain a significantly positive effect for formal membership. The same result obtains if we differentiate between types of membership and country groups.

1 Introduction

Any country becoming a member of the WTO is expected to honor its guiding principles. These are: a) Most-favored nation treatment, b) national treatment of foreign goods, services and intellectual property rights, c) multilateral negotiations on reciprocal reductions of trade barriers, d) fair competition rules related to dumping and subsidies, including a mechanism of dispute settlement, and e) preferential treatment of developing countries.[1] Judging from the degree of compliance, member countries in many instances appear to question the national advantage of adhering to these principles. Membership often seems to come at the cost of foregoing preferred policies, or having to yield concessions with unwelcome effects. Yet, overall these principles should lead to a transparent and predictable world trading environment that features open markets, thus enhancing world efficiency.[2] Therefore, WTO membership is commonly regarded as a key vehicle to enhance the growth and development perspectives of less developed countries.

In 1947, there were 23 founding signatories of the GATT. Presently, as many as 153 countries are members of the WTO. Moreover, the GATT/WTO was remarkably successful in reducing the level of trade barriers through 8 successive rounds of multilateral negotiations. On average, the import tariffs applied by GATT/WTO members have fallen to levels that are a mere quarter of what they were after the Second World War.[3] There was thus a widening as well as deepening of trade liberalization throughout the 6 decades that the GATT/WTO has been in existence. This was paralleled by an enormous post World War II increase in world trade, relative to world production. Between 1950 and 2005, the average annual growth rate of the volume of world exports was 6.2 percent (7.5 percent for manufactures), compared

[1] See the WTO's self-characterization on http://www.wto.org/, as well as chapter 3 in Bagwell & Stagier (2003).

[2] See Bagwell & Staiger (2003) for an in-depth analysis of the efficiency-enhancing potential of a WTO-like world trading system.

[3] The negotiations are about tariff *ceilings* (or *bindings*). *Applied* tariffs are typically below the ceilings. The WTO estimates that prior to the first round of negotiations (Geneva 1947), the average applied tariff rate across all member countries was between 20 and 30 percent. The Geneva round has brought a weighted US tariff reduction by 26 percent, with a cumulative further cut by 15 percent through the next four rounds of negotiation (Annecy 1949, Torquay 1951, Geneva 1956, Dillon 1962). Next came three more ambitious rounds of negotiations (Kennedy 1967, Tokyo 1979, and Uruguay 1994), each with average applied tariff cuts by well above 30 percent for industrial countries, affecting an ever larger amount of world trade. Thus, for the US, Canada and the major European countries, the *import-weighted average applied tariff rate* went down from 15 percent in 1952 to 4,1 percent in 2005. The negotiations also included reductions in quantitative restrictions and other non-tariff barriers, as well as export subsidies. For more details, see the WTO World Trade Report 2007. For the entire world, the World Bank estimates a reduction of the *unweighted average of import tariffs* tariff from 26.3 percent in 1986 to 8.8 in 2007; see the Data & Research Website of the World Bank.

with a real GDP growth of 3.8 percent.[4] The GATT/WTO almost routinely receives credit as a *causal* factor for this increase in world trade.

Somewhat surprisingly, however, when Rose (2004a) set out to quantify the trade-enhancing role of WTO membership in an econometric study of world trade based on the gravity equation, he ended up concluding that *"we currently do not have strong empirical evidence that the GATT/WTO has systematically played a strong role in encouraging trade"*. In a companion paper, Rose (2004b) has studied the trade policies pursued, to arrive at the conclusion that WTO member countries also do not follow more liberal trade policies than non-members. These papers have questioned the conventional view of the GATT/WTO as a significant trade-promoting institution.

However, subsequent literature has readdressed the issue, adding specific pieces of revisionist evidence. Thus, Subramanian & Wei (2007) have shown that WTO membership appears more valuable (in terms of trade creation) to industrial countries than developing countries, while Tomz, Goldstein & Rivers (2007) have emphasized that WTO membership does promote trade, if defined to include non-formal (or de facto) compliance with GATT/WTO rules. However, doubts remain. Rose (2007) points out certain puzzles that cast doubt on the trade-promoting role of non-formal membership. Eicher & Henn (2008) question the findings of Subramanian & Wei (2007) on the grounds of a more comprehensive treatment of preferential trade agreements alongside WTO membership, as well as on econometric grounds.[5]

This paper revisits the issue by looking at a place of potential evidence that has so far received relatively scant attention. By restricting his sample to country pairs where trade is strictly positive, Rose (2004a) has ignored the possibility that WTO membership may be important for whether or not two countries trade with each other at all. This is the so called *extensive* margin of world trade, as opposed to the *intensive* margin relating to how existing trading relationships evolve through larger or smaller quantities traded. Felbermayr & Kohler (2006) present detailed evidence on the relative importance of these two margins, concluding that the postwar increase of world trade took place through both, larger quantities traded (the intensive margin) and an increase in the number of country pairs that engage in trade (extensive margin). The question then is whether WTO membership comes out as a stronger trade promoting force if movements at the extensive margin of trade are adequately taken into account. Evidence pointing in this direction is presented in Liu (2007), as well as in Helpman, Melitz & Rubinstein (2008). In this paper, we present evidence that sheds new

[4]See again the most recent WTO World Trade Report 2007.

[5]See also Rose (2006) for a comprehensive reply to the literature subsequent to Rose (2004a).

light on this issue.

More specifically, we contribute to the literature as follows. First, we present a comprehensive review of the literature which clearly identifies the open issues. Secondly, we develop the simplest possible theoretical framework for a gravity equation that allows for country pairs with zero trade, thereby generalizing the well-known Anderson and van Wincoop (2003) model. Third, we estimate this model using a Poisson approach, in order to deal with the inherent non-linearity of any gravity model of bilateral trade that allows for zero trade. We argue on grounds of careful data inspection that any empirical analysis of the importance of the *extensive* margin should rely on cross-sectional variation, rather than panel estimation. And finally, when comparing the extensive margin and intensive margin effects of GATT/WTO membership, we also address the concerns raised in Subramanian & Wei (2007) about appropriate country pooling, as well as the issue of informal participation in the GATT/WTO raised by Tomz et al. (2008). Overall our results question the view that WTO membership has had much of an influence at the extensive margin of world trade.

The structure of the paper is as follows. In section 2, we first discuss the present state of the literature. In section 3, we present a simple theoretical model of the gravity equation that incorporates zero trade as an equilibrium outcome determined among other things by GATT/WTO membership. In section 4, we discuss our data base for econometric estimation, including some preliminary descriptive exploration that guides our estimation strategy. Section 5 takes a Probit-look at the extensive margin of trade, while section 6 presents results from a comprehensive estimation of the nonlinear gravity model. Section 7 will summarize and draw conclusions from our findings.

2 State of the literature

Given the aforementioned consensus view of the GATT/WTO, it is not surprising that Rose's (2004a) finding has caught a great deal of attention. It seems to cast doubt on the GATT/WTO as a "success story" that exemplifies the virtues of multilateral trade liberalization. But perhaps one should not be too surprised. It is well known that the GATT/WTO was only partly successful in delivering trade policies toward freer trade. There were sectoral exemptions, most notably in agriculture and textiles, and there were country exemptions as well. For instance, up until 1995 developing countries were facing little demand for liberalization when they became members. Moreover, member countries have partly undone negotiated tariff cuts by introducing non-tariff barriers. They have also made extensive use — sometimes abusively — of anti-dumping and safeguard provisions, as well as the WTO's dispute settlement mechanism, with disruptive effects on trade.

In a companion paper, Rose (2004b) has substantiated this concern by examining whether GATT/WTO member countries have systematically followed more liberal trade policies than non-members. His conclusion is that *"there is little evidence that membership in the GATT/WTO has actually liberalized trade policy"*. Hence, the lack of a significant and robust *trade effect* demonstrated in Rose (2004a) may simply reflect the lack of a liberalizing *trade policy effect* of the WTO. But this explanation is not entirely convincing, since it is obvious from the above characterization that WTO membership involves more than what is observable in terms of its members' trade policies. At any rate, such an explanation would still leave us with a troubling verdict on the GATT/WTO, whose primary mandate is to foster more liberal trade policies.

Several contributions have questioned that such a verdict is justified. For instance, Subramanian & Wei (2007) argue that there is a systematic pattern of asymmetry in this nexus of WTO membership and trade policies. They conduct a Rose-type empirical analysis, but looking at unidirectional trade (imports) rather than total trade flows between any two country pairs. More importantly, they allow for WTO membership to play a different role for developing and industrial countries. They find a strong positive trade volume effect for the latter, but not for the former. For instance, their preferred specification implies that WTO membership has on average increased industrial countries' bilateral imports by as much as 175 percent, while the estimated coefficients imply a much lower effect for developing countries. In their view, this reflects a policy asymmetry in that developing countries did not utilize their WTO membership toward trade liberalization, while developed countries typically did. This asymmetry, in turn, is due to differentiated treatment of these two groups of countries by the GATT/WTO; see above. This view receives empirical support in that Subramanian & Wei (2007) also find larger membership effects after the change to a more demanding stance vis a vis new members that took place in 1995. They also disaggregate along the sectoral dimension, to find positive membership effects for trade in liberalized manufacturing (for all countries) and for trade in non-liberalized manufacturing (for developed countries). Unsurprisingly, no positive effect was found for textiles, footwear and food. What are we to conclude from this exercise of disaggregation? While the results are certainly revealing, in our view they can hardly be interpreted as unequivocal support of a trade increasing effect of WTO membership. Rather, they provide a gravity-based documentation of the partial failure/success of the GATT/WTO. Allowing *all* data to speak up in a unified way, Rose's (2004a) finding that there is no robust positive WTO membership effect on trade remains upheld, even after Subramanian & Wei (2007); see also Rose (2006).[6]

[6]Some studies have taken the concern about excessive country pooling to the extreme by looking at individual countries. Thus, Lissovolik & Lissovolik (2006) have found Russian trade with WTO members to

But a negative verdict on the GATT/WTO is still unjustified according to a further criticism raised by Tomz et al. (2007). Rose's results might partly be due to the fact that WTO-type MFN treatment was sometimes also granted to non-members. In a similar vein, the attempt to secure WTO accession might have triggered more liberal trade policies ahead of formal membership. In Rose's empirical strategy, either of these two cases militate against a significant trade effect of WTO membership, provided that MFN treatment and pre-accession liberalization did in fact lead to more trade. Tomz et al. (2007) therefore suggest a broader view of WTO membership which includes *de facto* participation without *formal* membership. They point out that the WTO explicitly provides for such participation by member countries' colonies, as well through provisional membership status during a country's accession negotiation. Redoing Rose's analysis with extended WTO coding of the data (including non-membership participation), they find positive trade effects that are significant, both statistically and economically. For instance, the difference between bilateral trade volumes of non-member participants of the WTO and non-participants is 140 percent. These "revisionist" results make the GATT/WTO appear in a more favorable light than was the case in Rose (2004a), as well as Subramanian & Wei (2007). However, Rose (2007) points out that closer inspection reveals certain puzzles that continue to cast doubt on a trade-promoting effect of WTO membership. For instance, it seems difficult to imagine why non-membership participation in the WTO should have a larger effect than membership, and the bilateral trade effects found by Tomz et al. do not appear to show up in aggregate trade.

In statistical terms, the point raised by Tomz et al. (2007) is that Rose's (2004a) analysis involves a measurement error in WTO membership, leading to a downward bias in coefficient estimates, while Subramanian & Wei (2007) point to country heterogeneity among WTO members, measuring heterogeneity through the state of development. But there is additional *unobserved* country heterogeneity, also among non-members, which may give rise to biased coefficient estimates. Baier & Bergstrand (2007) have thoroughly addressed this issue with respect to free trade agreements (FTAs), or more generally preferential trading agreements (PTAs). In terms of our analysis, their argument is that there may be *unobserved* country characteristics that influence trade, which may at the same time be systematically correlated with WTO membership. Conventional estimation of the WTO membership effect would then suffer from an omitted variables bias. Baier & Bergstrand (2007) suggest using dyadic fixed effects in a panel data set to avoid a bias from such endogeneity. In addition to controlling for endogeneity, this would also control for the time-invariant component of

be less than would be expected given its "gravity position". The explanation offered by the authors can be interpreted as evidence in the spirit of Subramanian & Wei (2007). Evenett & Gage (2005) have found mixed effects for WTO accession of Angola, Bulgaria, Ecuador and Jordan.

unobserved multilateral trade resistance.[7] Their empirical study demonstrates that in much of the earlier literature endogeneity has indeed caused a downward bias in the estimated coefficients for the trade effects of PTAs. In principle, the same could hold true also for the estimated WTO membership effects, although they do not address this in their paper. It is interesting to note that Rose (2004a) did run estimations with country fixed effects, which have consistently generated larger membership effects than those without. However, he still regards these coefficients as small, relative to other effects; see Rose (2006).[8]

Generally, however, true *endogeneity* seems a less severe problem with WTO membership as such than with PTAs. An upward bias would arise, for instance, if there is some unobserved dyad-specific variable which is both, positively correlated with WTO membership (possibly even in a causal way) and bilateral trade. Baier & Bergstrand (2007) list several examples in this vein for PTAs. Basically, the endogeneity concern arises if certain country pairs are more natural trading partners than others, and are therefore more likely to reach a regional trading arrangement (in addition to trading more), for reasons other than those observed "on the right-hand-side". But this argument seems less convincing for the WTO which is a multilateral, not a regional, trading arrangement. Jointly entering a *multilateral* agreement like the GATT/WTO seems a somewhat odd response to being natural *bilateral* trading partners. In a similar vein, the notion that certain countries are more natural trading partners than others for the "whole world" (instead of bilaterally) seems far-fetched. However, even absent endogeneity of WTO membership, heterogeneity along the lines of "natural trading partners" or preferential trading arrangements might still involve de facto correlations, such that ignoring PTAs as an explanatory variable for bilateral trade will lead to an omitted variables bias. A priori, an upward bias seems more likely than a downward bias. The initial studies by Rose (2004a,2005) did include PTA-regressors, but still failed to deliver a robust WTO membership effect on trade.

From a broader perspective, the WTO status is but one element of the trading arrangement that governs bilateral trade. Adding PTAs and the Generalized System of Preferences (GSP), one obtains a rich pattern of possible arrangements. Some authors have argued in favor of a mutually exclusive coding of trading arrangements. This means classifying trading arrangements in such a way that any country pair belongs to one and only one arrange-

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[7]The notion of multilateral trade resistance was introduced into the gravity approach by Anderson & van Wincoop (2003). Their estimation involves nonlinear constraints. Importer- and exporter-country fixed effects, are an easier way to control for multilateral resistance. Subramanian & Wei (2007) do this as well.

[8]Rose (2006) also emphasizes the difference in interpretation between fixed effects ("within") estimation of WTO-coefficients and estimates that also use cross section evidence. "Within" estimation asks whether accession increases trade for a given country pair, while cross-section estimates compares across different country pairs.

ment. Bilateral trading arrangements relying only GATT/WTO membership would then be identified in isolation, and the corresponding coefficient could be interpreted as a “pure” WTO effect. Alternatively, PTA and WTO membership could be coded independently, as in the original studies by Rose (2004a,2005). As Subramanian & Wei (2007) have pointed out, if cumulative arrangement effects (e.g. from WTO membership plus PTAs) are additive, then independent and mutually exclusive coding should deliver the same effects. However, additivity might be questioned. For instance, one may argue that membership effects are hierarchical. For instance, Subramanian & Wei argue that a PTA dominates all WTO effects, meaning that WTO membership does not further increase bilateral trade if two countries already belong to a PTA. If this is the case, then identification of the true WTO effect requires hierarchical coding, whereby WTO membership is coded only for country pairs that do not also belong to the same PTA (and similarly for importer GSP-status). If the trading arrangements in fact work in a hierarchic manner, then independent coding would lead to an underestimation of the pure WTO effect. It would seem that a truly general estimation of arrangement effects requires mutually exclusive coding. Having estimates based on such data at hand, one can then use an additivity (or some other) assumption to calculate composite effects.

Eicher & Henn (2008) do this for the hierarchy assumption employed by Subramanian & Wei (2007), similarly distinguishing between industrial and developing country PTA effects. They corroborate the finding that the WTO works for industrial countries, but not for developing countries. Their composition exercise adds a further insight: Industrial countries find almost no additional PTA effect, while developing countries see an average PTA effect equal to 214 percent, which is even more than the WTO effect for developing countries. Eicher & Henn (2008) also distinguish between different types of PTAs, and they add bilateral fixed effects. This gives full control over all unobserved country-pair heterogeneity, in addition to controlling for the time-invariant component of multilateral resistance. Estimation of WTO and PTA effects is then based entirely on within-variation (time dimension). Their main message is that almost all significant WTO affects found by Subramanian & Wei (2007) and Tomz et al. (2007) vanish. This is a severe blow to the “revisionist” conclusions suggested by recent literature. Importantly for our purpose, however, Eicher & Henn (2008) restrict their data to positive trade pairs. The extensive margin of trade thus remains unexplored in their analysis.[9]

[9] It is worth pointing out here that “revisionist” conclusions have also been drawn based on non-parametric methods. Thus, Chang & Lee (2007) use non-parametric matching techniques to check for a trade effect of “WTO treatment”, coming up with large positive effects of WTO membership on trade. Ambiguity still seems to prevail.

This leads to what thus seems to be the ultimate line of defense for the WTO, which is at the core of this paper. Rose (2004a) and most of the subsequent literature estimate the gravity equation on data that exclude country pairs where bilateral trade is zero. As pointed out in the introduction, this is a potentially severe restriction. Intuitively, WTO membership may play a role, not only for how much two countries trade with each other, but also for whether they trade with each other at all. Technically, estimating a gravity equation on non-zero observations alone suffers from an omitted variable bias. Unfortunately, the standard theoretical underpinning of the gravity equation used in this type of literature does not allow for zero trade. Felbermayr & Kohler (2006) suggest an ad-hoc modification of the gravity model where zero trade emerges as a "corner-solution" in the sense of Wooldridge (2002). Estimating this model using Tobit techniques they have generally found the omitted variables bias to be empirically important.[10] They also find that including the extensive margin does make a difference for the estimated trade effect of WTO membership, generating more evidence for a positive trade effect from membership. Helpman et al. (2008) use a framework with heterogeneous firms to model the extensive margin, and they run a Heckman-type procedure for empirical estimation. Although they look at the WTO issue only in a peripheral way, they do find a significant effect of joint WTO membership in their Heckman selection equation. According to their estimate, the likelihood of positive trade increases by 15 percentage points if two countries belong to the WTO. Liu (2007) runs panel regressions, also including zero trade country pairs in his sample when looking at WTO membership effects. He finds strong WTO effects at the extensive margin of world trade.

This paper is thus not the first to explore WTO membership effects at the extensive margin of world trade. However, discussing the research that has followed his seminal contribution, Rose (2006, p. 18) concludes "*I'm now persuaded that membership in the GATT/WTO encourages the creation of trading links where none might otherwise exist. How important this is to world trade and welfare is currently unclear to me; I look forward to more work in the area.*" Our paper identifies and tries to fill several gaps that still exist in the literature.

Besides providing a concise yet thorough survey of the existing empirical literature, we offer the following three main contributions. First, we go back to theory in developing the simplest possible gravity model of bilateral trade that allows for zero trade as an equilibrium outcome. Although we call our model a "corner solutions" version of the gravity model, it goes well beyond Felbermayr & Kohler (2006), while at the same time being much simpler

[10] The empirical implementation of this approach in Felbermayr & Kohler (2006) has revealed that the increasing force of distance as a trade barrier, transpiring from recent gravity estimates of the distance coefficient ("distance puzzle"), can partly be explained as reflecting an omitted variables bias from ignoring the extensive margin.

than Helpman et al. (2008).[11] More specifically, we generalize the Anderson & van Wincoop (2003) model in the simplest possible way, in order to have the extensive margin of trade explicitly show up in the familiar multilateral resistance terms. Second, our theoretical model can be consistently estimated using the Poisson pseudo-maximum-likelihood procedure with exporter and importer fixed effects, as suggested by Santos & Tenreyro (2006). We work with averaged cross-sections, since within and between variation on the extensive margin is extremely noisy in the data. We compare results that are obtained with and without inclusion of zero trade flows into the sample in order to gauge the quantitative importance of those zero-trade observations. Third, we build on the Subramanian & Wei (2007) and Eicher & Henn (2008), as well as Tomz et al. (2007) and check whether their empirical strategies make a difference for the importance of the extensive margin. [12]

3 A simple model of the extensive margin

Empirical analysis of WTO membership at the extensive margin of world trade requires a suitable theoretical framework of world trade. We follow the literature in using the gravity model. However, we modify this model to incorporate the possibility of zero bilateral trade. The gravity equation for bilateral trade arises whenever demand satisfies the Inada conditions, and if in equilibrium any one good is produced in only one country. In a world of comparative advantage, this condition is met if equilibrium happens to involve complete specialization, as noted by Deardorff (1998). The preferred justification in the literature, however, relies on the monopolistic competition model of trade, enriched by features of geography, where complete specialization is a necessary outcome due to product differentiation. The most complete derivation of the gravity equation along these lines is Anderson & van Wincoop (2003). Importantly, this derivation features trade barriers only of the iceberg cost type, whence all goods are sold everywhere, albeit for different c.i.f. prices. It also features Dixit-Stiglitz preferences which satisfy the Inada conditions, hence there is positive trade between any pair of countries. In other words, there is no extensive margin of world trade.

[11] More details on the meaning of "corner solutions" will follow in section three.

[12] The paper most closely related to ours is Liu (2007) who similarly uses Poisson regressions to address the extensive margin. However, there are several key differences. First, our analysis is based on an explicit model of the extensive margin. Second, we use directed trade as the dependent variable, thus avoiding the silver medal mistake pointed out by Baldwin and Taglioni (2006), and we deliberately abstain from panel estimation. Third, we duly take into account third country effects through multilateral resistance. And fourth, we explore the above mentioned distinctions emphasized by Subramanian & Wei (2007) and Tomz et al. (2007). Since we do not exploit the time dimension of the data, our results are probably best viewed as complementary to those of Liu.

Empirically, there are a lot of country pairs where bilateral trade is zero.[13] Acknowledging the lack of an extensive margin in the underlying theory, existing literature mostly estimates the gravity equation on data that is restricted to country pairs with positive trade. This risks misspecification of the regression and ignores potentially important information. Exploiting this information requires a model allowing for zero trade. Felbermayr & Kohler (2006) suggest a simple ad-hoc modification of the gravity approach where zero trade cases arise as "corner-solutions", and which may be estimated using Tobit techniques. The modification rests on a threshold level of bilateral trade that needs to be reached for governments to incur the investment cost for *public* infrastructure as required to support trade. Helpman et al. (2008) develop a model of monopolistic competition with *private* fixed cost of serving foreign markets. Their model combines such cost with firm heterogeneity in marginal cost, to generate two extensive margins of trade. There is an extensive *country* margin separating zero trade country pairs from trading pairs.[14] In addition, for each country pair the model also generates an extensive *firm* margin that separates firms exporting to the partner country from those who don't. The model is in the spirit of Melitz (2003), although it is partial equilibrium in that it does not incorporate a resource constraint, with an exogenous mass of firms. There is multiple selection of *existing* firms into potential export markets. The model is geared toward estimation of a generalized gravity approach including the two extensive margins, using a two-step estimation procedure a la Heckman (1979).

For the purpose of this paper and for many other applications, a much simpler model without firm heterogeneity suffices to motivate the empirical strategy. On the demand side, we assume Dixit-Stiglitz-type preferences, identical for all countries, with a constant elasticity of substitution $\sigma > 1$ between different varieties of goods. Preferences are fully symmetric across all potential varieties, independent on the country of origin. Using p^{ij} to denote the c.i.f.-price in country j for a variety arriving from country i, demand for this variety may be written as $D^{ji} = A^j \left(p^{ij}\right)^{-\sigma}$, where $A^j := Y^j \left(P^j\right)^{\sigma-1}$. In this expression, Y^j is equal to country j's GDP, and P^j is the exact price index (unit-expenditure function), depending on

[13] As indicated above, Felbermayr & Kohler (2006) present a decomposition of the postwar evolution of world trade into the extensive and the intensive margin.

[14] An extensive country margin of world trade also arises in Eaton & Kortum (2002) who employ a stochastic representation of the pattern of (Ricardian) comparative advantage among many countries with perfect competition. Adding geography in the form of trading cost, they arrive at a gravity-type equation for bilateral trade. The key difference to the standard gravity model is that for any good all countries are perfect substitutes as sources of supply. In their model, two countries end up with zero bilateral trade, if for the entire range of goods each of them finds cheaper supply from third countries, due to their idiosyncratic technology and geography.

prices of all varieties shipped to market j.[15]

There are $C+1$ countries, and we use i to indicate a country of production. Each firm produces its own variety, using a well defined efficiency unit of a bundle of inputs, with associated minimum unit-cost c^i, which is specific to the country where a firm is located. Variation in c^i will be explained below by varying factor endowments and/or varying overall productivity across countries. Unlike Helpman et al. (2008), we assume that all firms have the same productivity in terms of both marginal and fixed cost. We use a to denote the constant marginal input requirement, while f denotes the fixed cost of production, all in terms of efficiency units of the input bundle.

Serving a foreign market entails two types of cost. One is variable trade costs, assumed to be of the iceberg-type and captured by a parameter $\tau^{ij} > 1$, where i and j indicate the sending and receiving country, respectively. In addition, there is a fixed cost f^{ij} that each firm located in country i has to bear when entering an export market j.[16] Fixed costs are in terms of the input bundle with minimum unit cost c^i. Domestic sales do not require any of these costs, whence $\tau^{ii} = 1$ and $f^{ii} = 0$. Variable and fixed trade cost are defined to include both, natural as well as policy-induced factors of trade resistance.

Profit maximization by a representative firm located in country i, assuming any P^j to be given, implies a markup price equal to

$$p^{ij} = \tau^{ij} c^i a/\rho, \tag{1}$$

where $\rho := (\sigma-1)/\sigma$ from the Dixit-Stiglitz preferences. The price for domestic sales is equal to c^i/ρ. A typical firm locating in country i will perceive maximum profits to be earned on exports to country j equal to

$$\pi^{ij} = (1-\rho)\rho^{\sigma-1} A^j (c^i a \tau^{ij})^{1-\sigma} - c^i f^{ij}. \tag{2}$$

Obviously, the firm will choose to export to country j only if $\pi^{ij} \geq 0$. This condition may be rewritten as

$$(1-\rho)\rho^{\sigma-1}(c^i a)^{1-\sigma} \geq f^{ij}(\tau^{ij})^{\sigma-1} / A^j. \tag{3}$$

We can now envisage all potential trading partners of country i as being ranked, such that $f^{ij}(\tau^{ij})^{\sigma-1}/A^j$ falls monotonically, as $j \neq i$ increases from 1 up to C. Note that this ranking is specific to the exporting country i. Then, there will be a marginal country j^i, such that for all countries $j \leq j^i$ condition (3) is satisfied, while for all countries $j > j^i$ it is violated.

[15] GDP replaces the level of expenditure, assuming balanced trade.

[16] The model bears some resemblance to Schmidt & Yu (2001) who introduce firm heterogeneity in fixed costs of exporting into a single export market.

This is the *extensive country margin* of exports for country i, determined independently for each export market. We introduce an integer-valued function

$$j^i = j^i(c^i, \boldsymbol{\tau}^i, \mathbf{f}^i) \tag{4}$$

to indicate that country i's extensive margin of trading partners j^i depends on its entire pattern of iceberg- and fixed trade resistance, which appear in vector forms $\boldsymbol{\tau}^i$ and $\mathbf{f}^i$ in equation (4). We use J^i to denote the index set of all $j \leq j^i$. Thus, country i will have positive exports only to countries $j \in J^i$. Notice that even if trade costs are symmetric in both directions, trade flows may be unidirectional, since c^i is country specific, and the trade resistance terms need not be symmetric in terms of direction. It should be noticed that the key "relative price" determining profitability of bilateral exports is $(c^i \tau^{ij} a)/P^j$, in addition to the size of the market relative to fixed cost, Y^j/f^{ij}.

We now introduce a latent variable $\rho^{\sigma-1} A^j (c^i \tau^{ij})^{1-\sigma}$, which is the potential value of exports to country j per firm located in country i. The actual per firm value of exports from i to j then is

$$p^{ij} x^{ij} = \begin{cases} \rho^{\sigma-1}(c^i a \tau^{ij})^{1-\sigma} A^j, & \text{if} \quad i \in J^i \\ 0 & \text{otherwise} \end{cases} \tag{5}$$

This is the "corner-solutions" formulation of bilateral exports, where $c^i f^{ij}/(1-\rho)$ is a threshold value for the *latent* variable $\rho^{\sigma-1}(c^i a \tau^{ij})^{1-\sigma} A^j$ that determines zero "corner-solutions" of bilateral exports to all countries $j > j^i$ (or, equivalently, $j \notin J^i$). Aggregate bilateral exports are $X^{ij} = N^i p^{ij} x^{ij}$, evaluated at c.i.f. prices.[17]

We now introduce two aggregate measures of the export markets served by a firm located in country i:

$$\Theta^i := \frac{1}{A^i} \sum\nolimits_{j \in J^i} A^j (\tau^{ij})^{1-\sigma} \quad \text{and} \quad \digamma^i := \sum\nolimits_{j \in J^i} f^{ij}. \tag{6}$$

The variable Θ^i the "aggregate size" of foreign markets reached by a representative firm of country i, relative to country i's domestic market, using A^j to measure market size, and scaling down market size by the iceberg trade cost term $(\tau^{ij})^{1-\sigma}$. It is closely related to what is sometimes called the nominal export market potential of country i. More precisely, Θ^i is the ratio of c.i.f. export sales to domestic sales by a representative firm located in country i. In what follows domestic sales are indicated by $p^{i0} x^{i0}$. The variable $\digamma^i$ is a simple measure of the entire fixed costs of exporting incurred by each firm in country i, again measured in

[17]In this paper, we use the term "corner solutions" model for any model that involves zero trade as an equilibrium outcome of the same mechanisms that also determine the volume of trade. Any such model involves a non-linear relationship between trade (allowing for zeros) and the covariates. Wooldridge (2002) uses the term "corner solutions" to describe Tobit estimation techniques to incorporate the nonlinearity. In out empirical analysis, we rely on a Poisson pseudo maximum likelihood estimator; see below.

efficiency units of the input bundle with a per unit cost equal to c^i. Both of these measures are determined by the extensive country margin of exports which is determined as described above. For given trade costs, both Θ^i and $\digamma^i$ are increasing in $j^i(\boldsymbol{\tau}^i, \mathbf{f}^i)$.

It is important to make a distinction between the impact of infra-marginal trade liberalization which operates through τ^{ij} and f^{ij} for $j \in J^i$ with a given j^i, and the impact of trade liberalization at the margin, which increases j^i through either a reduction in τ^{ij} or f^{ij} for some $j > j^i$. However, we first continue describing the equilibrium for a given j^i.

Profits on actual exports from i to $j \leq j^i$ are equal to $[(c^i a\tau^{ij}/\rho) - c^i a\tau^{ij}]x^{ij} - c^i f^{ij}$. Moreover, denoting domestic sales by a representative country-i-firm by x^{i0}, market clearing, $D^{ji} = x^{ij}$, requires $x^{ij} = (A^j/A^i)(\tau^{ij})^{-\sigma}x^{i0}$ and $p^{ij}x^{ij} = (A^j/A^i)(\tau^{ij})^{1-\sigma}p^{i0}x^{i0}$, since $p^{ij} = \tau^{ij}p^{i0}$. Assuming free entry, we may then invoke a zero profit condition of the form

$$x^{i0}a\frac{1-\rho}{\rho}\left(1+\Theta^i\right) - \left(\digamma^i + f\right) = 0. \tag{7}$$

This determines the level of domestic sales per firm in a familiar way, except for the appearance of two measures of the extensive country margin of exports, $\digamma^i$ and Θ^i. Total output and sales per firm are determined solely by market potential and fixed costs, $x^{i0} = (\sigma - 1)\left(\digamma^i + f\right) / \left[a\left(1+\Theta^i\right)\right]$. Note also that we have the GNP-identity $N^i x^{i0} c^i a\left(1+\Theta^i\right)/\rho \equiv Y$, hence

$$p^{i0}x^{i0} = \frac{Y^i}{N^i\left(1+\Theta^i\right).} \tag{8}$$

We now close the model by introducing endowments and factor market clearing. Using the scalar V^i to denote country i's endowment with the input bundle (in efficiency units), the equilibrium number of firms emerges from the full employment condition which is written as $N^i\left[aX^{i0}(1+\Theta^i) + f + \digamma^i\right] = V^i$. In view of (7), this simplifies to

$$N^i = \frac{V^i}{\left(\digamma^i + f\right)\sigma}. \tag{9}$$

Thus, the equilibrium number of firms is unaffected by Θ^i, but it falls with an increase in $\digamma^i + f$, relative to the resource base V^i. On the other hand, the equilibrium volume of domestic sales per firm is independent on the resource base V^i. These are familiar properties of the Dixit-Stiglitz version of the monopolistic competition model.

Finally, factor prices are determined such that the cost-minimizing input levels per efficiency unit of the input bundle add up to the country's endowment for each of the factors present. More specifically, assume that $c(w)$ is a dual characterization of the technology, common to all countries, according to which K different factors may be combined to generate an efficiency unit of the input bundle required in the variable and fixed cost of production and trade, with w denoting the vector of factor prices.[18] Then, $e^i c_k(w^i)$ gives the cost-minimizing *physical* quantity of factor k used to generate an *efficiency unit* of the input bundle according to country i's (Ricardian) level of technology, captured by the efficiency parameter e^i. Factor market equilibrium in country i then requires

$$e^i c_k(w) = V_k^i / V^i \quad k = 1 \dots K \tag{10}$$

for all $k = 1 \dots K$, where V_k^i denotes the physical endowment of country i with factor k. The factor cost advantage of country i introduced above then emerges as $c^i = e^i c(w^i)$, where w^i is the factor price vector satisfying the aforementioned factor market equilibrium condition.

We are now ready to derive a gravity equation for bilateral trade. Aggregate bilateral exports X^{ij} from i to j are given by $N^i p^{ij} x^{ij} = N^i (A^j / A^i)(\tau^{ij})^{1-\sigma} p^{i0} x^{i0}$. Replacing for $p^{i0} x^{i0}$ from (8), we have

$$X^{ij} = \begin{cases} Y^i Y^j \frac{(\tau^{ij}/P^j)^{1-\sigma}}{\sum_{j=0}^{j^i(c^i, \boldsymbol{\tau}^i, \mathbf{f}^i)} A^j (\tau^{ij})^{1-\sigma}} & \text{if } j \in J^i \\ 0 & \text{otherwise} \end{cases} \tag{11}$$

Note that the denominator in the first line is equal to $1 + \Theta^i$. Remembering that $A^j := Y^j (P^j)^{\sigma - 1}$, this leads to a generalized Anderson & van Wincoop (2003) representation of the gravity approach.[19] The key difference to note here is that the denominator in the first line involves the extensive country margin $j^i(c^i, \boldsymbol{\tau}^i, \mathbf{f}^i)$. The full Anderson & van Wincoop system emerges, if we recognize that the importing country's true price index P^j in the numerator of (11) has c.i.f. prices p^{ij} related to the respective home prices p^{i0}, now holding j fixed and varying the country of origin i. This leads to a second type of extensive margin where exporting countries select themselves into exporters and non-exporters for any importing country j. We may indicate this margin through an index set I^j whereby $i \in I^j$ iff $j \in J^i$.

We do not want to impose a symmetry assumption that would allow us to derive the full Anderson & van Wincoop system. The general idea of our approach is easy enough to see from (11) above. The essential points are as follows. First, the "gravity term" on the right-hand side of (11) is a latent variable which is strictly positive for all country pairs. However, firms in country i expect negative profits to be earned on exports to any country $j \notin J^i$, due to fixed costs f^{ij}, and observed exports X^{ij} will, therefore, be zero. These are the "corner-solutions". Note that by assumption firm decisions are made taking all aggregate variables, i.e., all price indices and multilateral resistance terms (involving extensive country margins

[18]The model actually does not depend on the assumption of a uniform technology; we could allow for country-specific minimum cost function $c^i(w)$.

[19]See also Appendix II in Helpman et al. (2008).

j^i), as given. The number of firms in each exporting country is endogenously determined through (9).

The second point to observe relates to the WTO membership effect on trade. There is a *direct* trade promoting effect at the intensive margin through a lower τ^{ij}, if countries i and $j \in J^i$ are *both* members of the WTO. According to (11), these countries should have more bilateral trade than country pairs that are otherwise similar, but where both are outside the WTO or only one is a member. There is a perfectly analogous effect at the extensive margin, making positive trade for member country pair a more likely event than for a non-member pair. This may operate through lower τ^{ij} or lower f^{ij} in the denominator of (11). Notice that f^{ij} plays no role for the intensive margin effects of WTO membership, but may be important for extensive margin effects.

In addition, there are *indirect* ("third-country") effects that operate through the multilateral resistance channel. Suppose, for instance, that two countries i and $j \in J^i$ are *both* members of the WTO, and there is a third country $k \notin J^i$ with no exports from i to k. Now suppose that k joins the WTO and through lower τ^{ik} and or lower f^{ik} it jumps the extensive export margin of country i, i.e., k moves into J^i in the upper limit of the summation in (11). Bilateral trade resistance between i and j, *relative* to these two countries' multilateral resistance has *increased*, assuming that τ^{ij} as such remains unchanged. As a result, exports from i to j fall through an extensive margin effect of country k's WTO accession. The same effect will not be observed for *non-member* country pairs, but will be observed also if j is not in the WTO. Any study that looks only at the intensive margin and ignores the extensive margin multilateral resistance effect will thus underestimate the trade effect of WTO membership.[20] Intuitively, what we have here is something like a trade diversion effect that comes about through the resource constraint of the exporting country. Newly established trade between i and k draws away resources from exports to existing trading partners.

4 An empirical model

Our ultimate goal is to empirically quantify the effect of WTO membership. Our emphasis lies on an appropriate treatment of the extensive margin of trade, based on our corner solutions model of the gravity model developed above. While that model does highlight the possibility of zero trade between certain country pairs, it does not yet offer a workable estimation equation. A key feature of the "corner-solutions" gravity equation is that it is *non-linear*.

[20]We have seen in the preceding section that allowing for multilateral resistance through fixed effects has indeed tended to increase the estimated membership effect.

A possible estimation approach to such an equation is to rely on Tobit techniques, as in Felbermayr & Kohler (2006). This paper uses a different approach, recently suggested by Santos & Tenreyro (2006), that has a number of advantages over the Tobit approach.

According to this approach, the empirical model based on equation (11) is as follows. Recognizing that X^{ij} can be zero, we follow Santos and Tenreyro (2006) and write equation (11) as an exponential model

$$X^{ij} = \exp\left[(1-\sigma)\ln\tau^{ij} + K^i + K^j\right], \tag{12}$$

where $K^i := \ln Y^i - \ln\left[\sum_{j=0}^{j^i(c^i,\boldsymbol{\tau}^i,\mathbf{f}^i)} A^j (\tau^{ij})^{1-\sigma}\right]$ and $K^j := \ln\left[Y^j \left(P^j\right)^{\sigma-1}\right]$. Note that i and j denotes the exporter and importer country, respectively. The estimation strategy proposed by Anderson & van Wincoop (2003) or Feenstra (2004) amounts to using an array of exporter- and importer-specific dummy variables to control for K^i and K^j. This approach perfectly controls for the multilateral resistance term, which would otherwise have to be simulated and which cannot be directly measured in the data. Failing to control for the multilateral-resistance term is likely to cause severely biased estimates, as Anderson and van Wincoop (2003) have shown. It should also be noticed that the use of dummy variables does away with the awkward question of proper discounting of trade and GDP values; see Baldwin & Taglioni (2006). It allows estimation including observations of country pairs where GDP data is not available (or unreliable) for either the importer or the exporter. Importantly, the country-dummies are perfect controls for country-specific policies that apply to all trading partners, regardless of WTO membership. Many non-tariff and technical barriers to trade have this characteristic.

Typically, researchers add a multiplicative error term ε^{ij} to (12), take logs and substitute dummy variable vectors $\boldsymbol{\nu}^i$ and $\boldsymbol{\nu}^j$ for K^i and K^j, respectively. This yields the familiar log-log gravity equation $\ln X^{ij} = \alpha \ln\tau^{ij} + \boldsymbol{\beta}^j\boldsymbol{\nu}^j + \boldsymbol{\beta}^i\boldsymbol{\nu}^i + \ln\varepsilon^{ij}$, where $\alpha := 1-\sigma$ is interpreted as the trade-cost elasticity of bilateral exports and $\boldsymbol{\beta}^i$ and $\boldsymbol{\beta}^j$ are vectors of parameters associated with the exporter and importer country dummies introduced to control for K^i and K^j, respectively. In a typical empirical gravity equation (e.g., Anderson & van Wincoop, 2003), real trade costs τ^{ij} are specified as some multiplicative function of policy-induced trade barriers and a host of geographic variables $\tau^{ji} = \left(1+t^{ij}\right)\cdot\left(\mathrm{DIST}^{ij}\right)^{\delta}\cdot\exp\left[\left(1-\mathbf{CAT}^{ji}\right)\boldsymbol{\gamma}\right]$, where $(1-\mathbf{CAT}^{ji})$ is a row vector of ones minus relevant categorical variables (with associated coefficients in $\boldsymbol{\gamma}$) affecting real trade costs τ^{ij}, in addition to distance DIST^{ij} and ad valorem tariffs t^{ij}. Notice that, as usual, coefficients may be interpreted as elasticities.

The problem with this procedure, as pointed out by Silva & Tenreyro (2006), is that unless the *variance* of ε^{ij} is independent on the variables $\mathbf{Z}^{ij} := \left\{\tau^{ij}, \boldsymbol{\nu}^i, \boldsymbol{\nu}^j\right\}$, the *expectation* of $\ln\varepsilon^{ij}$ will depend on these same regressors, leading to inconsistent OLS estimates. Moreover, taking logs generates "missing values" if for some country pairs bilateral trade is zero, as in the "corners solutions" above. This, in turn, may bias estimates, since the data may no longer be viewed as randomly sampled.[21]

Santos & Tenreyro (2006) suggest an approach that avoids these problems. Given that the log of a stochastic variable also depends on its variance, the estimation should be guided by the assumed relationship between $E\left[X^{ji}|\mathbf{Z}^{ij}\right]$ and $V\left[X^{ji}|\mathbf{Z}^{ij}\right]$, where $\mathbf{Z}^{ij}$ denotes the entire vector of explanatory variables. For want of a more specific information on this relationship a reasonable hypothesis might be that conditional variance of M^{ji} is proportional to its conditional mean, i.e. $E\left[X^{ji}|\mathbf{Z}^{ij}\right] \propto V\left[X^{ji}|\mathbf{Z}^{ij}\right]$. Santos and Tenreyro show that (12) can then be estimated by solving the following set of first order conditions

$$\sum\nolimits_{r=1}^{C(C-1)} \left[M^r - \exp\left(\mathbf{Z}^r\hat{\boldsymbol{\beta}}\right)\right] z_h^r = 0 \quad h = 1 \ldots H \tag{13}$$

where r indexes bilateral import relationships (ji). In this estimation criterion, $\mathbf{Z}^r$ denotes an $H \times 1$-dimensional vector of covariate observations (with element z_h^r), and $\hat{\boldsymbol{\beta}}$ denotes the corresponding vector of of parameter estimates. This estimator is equivalent to the Poisson pseudo-maximum likelihood (PPML) estimator. It is consistent under the correct specification of the conditional mean. Importantly, the data need not be Poisson[22], and X^r need not be integer-valued. Santos & Tenreyro (2006) implement this estimator for a cross-sectional gravity equation. They compare what they call "a traditional gravity equation" that does not properly account for multilateral resistance with an "Anderson- van Wincoop gravity equation", using country-specific dummy variables. They find that the inclusion of those dummies affects results quite substantially, in particular concerning categorical variables, such as WTO membership.

The key contributions to the literature on the trade effects of GATT/WTO membership mostly use OLS on pooled cross-sections. However, as emphasized above, the "corner-solutions" gravity model that allows for zero trade pairs is necessarily non-linear. Hence, the PPML approach suggested by Santos & Tenreyro (2006) is a much better way to estimate this model. The concerns raised in the recent literature surveyed in section 3 above can also be addressed using this approach. Indeed, this is the key contribution of the present paper.

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[21] Tobit and Heckman-type procedures can deal with the corner-solutions nature of equation (11); they are, however, not robust to misspecification of the error term.

[22] A maximum-likelihood estimator is called a pseudo maximum-likelihood estimator if it remains consistent even if the likelihood function is misspecified (Winkelmann, 2003).

5 Data and econometric strategy

Before turning to our estimation results, we briefly describe the estimation strategy that we believe is appropriate for our data set. We start with a brief description of the data.

Trade: Our trade data is from the May 2006 CD-ROM of the IMF's Direction of Trade Statistics (DoTS). Some of the literature (e.g., Rose, 2004) estimates the gravity equation on gross trade (exports plus imports). In line with theory, our model has bilateral exports as the dependent variable. However, we estimate the gravity equation on bilateral import data, measured in c.i.f. terms, as this data is usually deemed of higher quality. With the number of countries equal to $C = 181$ for the year 2000, we thus have $C(C-1) = 32,580$ bilateral trade relationships. Trade values are missing for 20.7 percent of these country pairs. The IMF codes 63.5 percent of the non-missing observations as zeros. Figure 1 provides an illustration of missing, zero, and positive trading relationships for the years 1948-2004. The data exhibit considerable discrete jumps in the share of missing observations, due to decolonialization or due to the break-up of countries, such as the Soviet Union in the 1990s. However, as a general trend, the share of active trading relationships has increased strongly over time, both measured in terms of total potential relationships and non-missing ones.

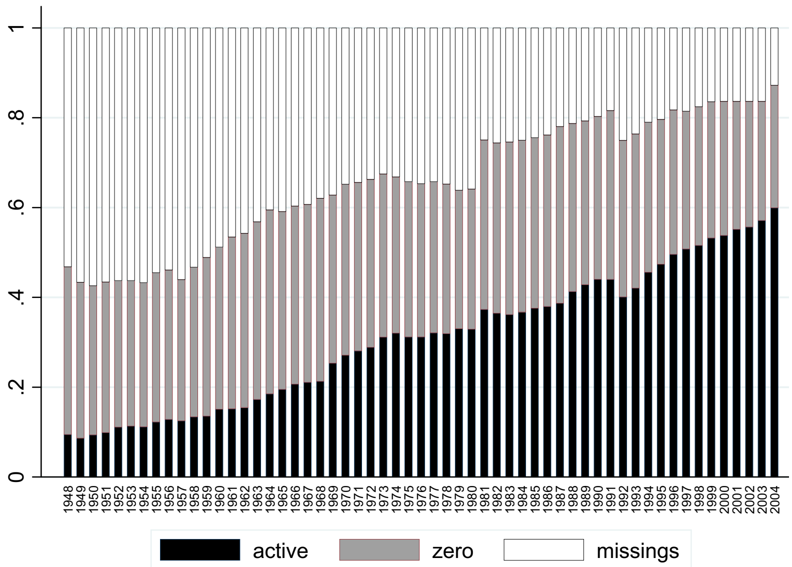

Figure 1: The relative importance of the extensive margin, 1948-2004.

One must be cautious, however, in interpreting variations at the extensive margin of trade cross time. For example, many country pairs exhibit frequent switches between active trade, zero trade, and missing. For example, between 1948 and 2004 we observe as many as 26 such switches for the Malta-Paraguay country pair; similarly for Sri Lanka-Tunesia. For 5 percent of all country pairs, switches occur more than 12 times. These frequent switchers are usually small, poor countries. Exploring the likelihood of switches across time, we find that a Probit model explaining zero trade does equally well in predicting the occurrence of missings. Hence, it seems unwise to focus too much on time-variance, at least in an analysis that explicitly focuses on the role of the extensive margin.[23] In our econometric analysis, we therefore take averages over certain time spans. More specifically, we define the following sub-periods characterized by certain stages of the GATT: Pre-Kennedy-round (1948-1967), pre-Tokyo-round (1968-1979), pre-Uruguay-round (1980-1994) and, finally, the (post-Uruguay) WTO-period (1995-2004). The first period covers a sequence of tariff-cutting rounds. The second period is characterized by a broadening of the scope of the GATT to include anti-dumping measures. The third period additionally involves rules on non-tariff measures and the so-called framework agreements. And the fourth period is marked by the transition from the GATT to the WTO.

Within these time intervals, missing bilateral imports are interpreted as zeros whenever at least one year features a non-missing observation. This procedure should reduce measurement error since it smoothens out idiosyncratic switches between zero, positive trade, and missings, although the strategy does reduce the amount of zero-trade observations, relative to a year-by-year treatment of the data. Hence, a corner solution of zero trade is defined as a case where, subject to the above treatment of missings, we observe zero trade for each year of the time respective time period. The summary statistics presented in Table 1 show that the share of active trading relationships has grown from about 48 percent in the pre-Kennedy period to 87 percent in the post-Uruguay times.

GATT/WTO membership: Tomz et al. (2007) emphasize the distinction between *formal* membership at the GATT or the WTO/WO, and *informal* participation in the institution, for example of colonies through their metropolitan colonizers, or through de facto compliance during negotiation ahead of formal membership. They find *factual* membership (i.e., compliance with GATT rules through either formal or informal participation) is significantly related to higher trade, while participation defined as formal membership does not.[24] In the subsequent econometric analysis, we want to see whether this distinction continues to matter when the analysis covers the extensive margin of trade and a theory-consistent estimation

[23]Indeed, comparing the raw data from COMTRADE with the DoTS reveals that the latter often has zeros and missings where the COMTRADE has small values (e.g., 1,000,000 dollar).

[24]Throughout this paper factual membership means compliance with GATT rules either through formal membership or de facto (informal) compliance as defined in Tomz et al. (2007).

Table 1: Summary Statistics

Variable	Pre-Kennedy N=13,155 Mean	Std. Dev.	Kennedy-Tokyo N=14,364 Mean	Std. Dev.	Tokyo-Uruguay N=20,388 Mean	Std. Dev.	Post-Uruguay N=22,516 Mean	Std. Dev.
I(M>0) indicator variable	**0.48**	**0.50**	**0.70**	**0.46**	**0.79**	**0.40**	**0.87**	**0.34**
ln Imports	13.20	2.77	13.30	3.80	13.49	4.12	13.79	4.27
Imports (USD mn)	7.75	92.60	45.30	497.00	123.00	1540.00	254.00	3200.00
WTO (formal)	0.14	0.26	0.35	0.47	0.39	0.45	0.38	0.29
WTO (formal) x IND	0.05	0.18	0.11	0.32	0.10	0.29	0.08	0.21
WTO (formal) x DEV	0.09	0.22	0.23	0.41	0.29	0.42	0.30	0.30
WTO (factual)	**0.41**	**0.43**	**0.58**	**0.48**	**0.57**	**0.47**	**0.41**	**0.28**
WTO (factual) x DEV	0.12	0.29	0.17	0.37	0.13	0.33	0.08	0.21
WTO (factual) x IND	0.29	0.42	0.41	0.48	0.45	0.48	0.32	0.30
Regional trade agreement (dummy)	0.00	0.01	0.00	0.04	0.01	0.07	0.00	0.05
GSP	0.00	0.01	0.15	0.29	0.17	0.37	0.08	0.19
ln Distance	8.72	0.77	8.72	0.77	8.69	0.79	8.69	0.79
Contiguity (dummy)	0.02	0.15	0.02	0.15	0.02	0.14	0.02	0.14
Common language (dummy)	0.18	0.39	0.18	0.38	0.15	0.36	0.14	0.35
Every colony (dummy)	0.02	0.13	0.02	0.13	0.01	0.12	0.01	0.12
Common colonizer (dummy)	0.12	0.32	0.12	0.32	0.11	0.31	0.10	0.31
Currently colony (dummy)	0.00	0.02	0.00	0.02	0.00	0.02	0.00	0.02
Same country (dummy)	0.01	0.10	0.01	0.10	0.01	0.10	0.01	0.10
Strict currency union (dummy)	0.00	0.05	0.01	0.06	0.00	0.05	0.00	0.03

approach is chosen.[25]

Figure 2 shows that the share of countries with *formal* membership in the GATT/WTO increases through time; it was less than 20 percent in 1948, rising to about 73 percent in the year 2001 when no more than 11 countries are classified as informal members.[26] Somewhat surprisingly, however, from the era of decolonialization (mid sixties) onwards, the share of countries *factually* participating in the GATT/WTO, has remained fairly stable. A similar picture emerges from the summary statistics in Table 1.

Other covariates: Our control variables are identical to those typically used in the literature; the data are from Rose (2004a). It includes dummies for joint membership in a regional trade agreement or a strict currency union and a dummy for whether the importer grants GSP (generalized system of preferences) status to the exporter. These variables may vary over time and represent trade policy controls. The model also includes geographical or cultural variables, such as geographical distance, contiguity, the existence of a common language, and a host of dummies reflecting the colonial relationship between an importer and the exporter. Rose's (2004a) estimation relies on data that covers the period 1948-1999.

[25]We use the classification of countries by formal and informal membership status, respectively, as provided on the website of Michael Tomz.

[26]The year 2001 is the latest period covered in the classification supplied by Tomz.

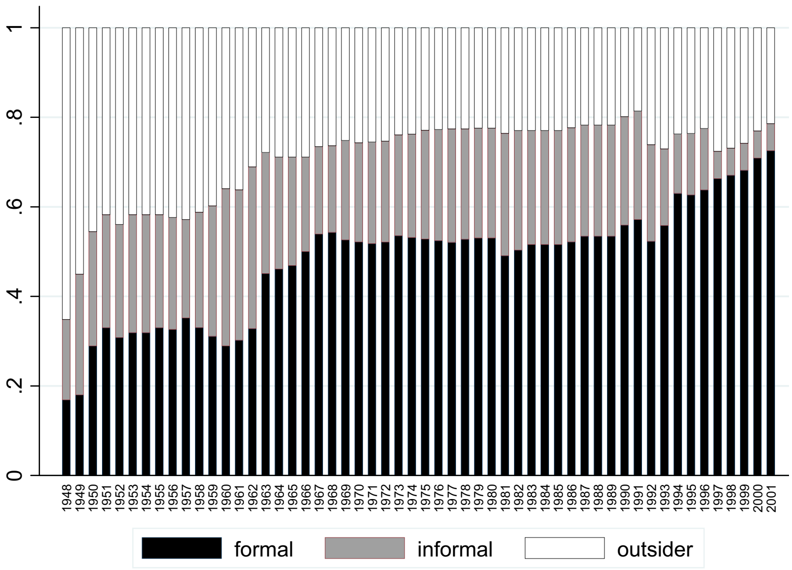

Figure 2: Formal / informal GATT/WTO members and outsiders, 1948-2004

We have updated our data to the year 2004 using information about WTO membership and regional trade agreements provided on the website of the WTO.

Econometric strategy: Our strategy differs from existing literature in several respects. First, unlike Rose (2004a), Tomz et al. (2007), Subramanian & Wei (2007) and Eicher & Henn (2008), we place emphasis in the extensive margin of trade. In a recent paper that looks at the extensive margin, Liu (2007) employs a panel framework, controlling for time-invariant bilateral effects (e.g., unobserved *bilateral* components of trade policy). He thus draws exclusively on within-variance (switches from outsider into member status over time). We have argued above that a close inspection of the data leads us to question the reliability of time series variation at the extensive margin of trade. We therefore propose an econometric strategy that relies on cross section variation rather than a panel framework. In that sense, our results are best viewed as complementary to Liu's. Our strategy has the further advantage that it allows us to trace the behavior of the WTO effect over time. It also avoids having to deal with the fact that over time new countries have been created through decolonialization and the break-up of the Soviet Union.[27]

We proceed in two steps. First, we look at the extensive margin in isolation, employing a *cross-section Probit* estimation framework for the aforementioned sub-periods representing characteristic episodes in the history of the GATT/WTO. In doing so, we also explore the differential role of formal vs. informal membership, as well as the difference between de-

[27] In Felbermayr & Kohler (2006), this is called the "pseudo-extensive" margin.

veloping and industrial countries. The second step then estimates the full corner-solutions gravity model, relying on Poisson pseudo-maximum-likelihood (PPML) procedure, in order to capture the non-linearity implied by corner solutions; see above. We avoid what Baldwin & Taglioni (2006) call the "gold medal mistake" by consistently using importer- and exporter fixed effects to control for unobserved variables.[28] Moreover, our specification is theory-grounded in the sense that we use bilateral imports, rather than total trade (exports plus imports) as our dependent variable. This also increases the number of observations for each period; see Table 1 for details.[29]

6 Estimation results

6.1 A Probit view on the extensive margin

In this section we look at role of WTO membership at the extensive margin of world trade independently of the intensive margin. We represent Probit regression results of a cross-section analysis in 4 different periods of GATT/WTO history. The dependent variable is an indicator variable I^{ij} that takes the value 1 if country i has strictly positive imports from country j, i.e., if $X^{ij} > 0$, and the value of zero otherwise. We present results from the Probit model, but logistic and the linear probability model give comparable results. All our regressions include country-effects, but they are not shown to save space. We decompose the effect of WTO membership according to whether the importer country is a developing or an industrial country. Using the country classification of Subramanian & Wei (2007), we define a dummy IND_i, such that $IND_i = 1$ if country i is industrialized and $IND_i = 0$ otherwise. We compute the interactions $WTO_{ij} \times IND_i$ and $WTO_{ij} \times (1 - IND_i)$ and jointly use them in our regressions, where $WTO_{ij} = 1$ indicates that countries i and j are both members of the WTO. For the sake of comparison, the variables included in the regression are the same as those by Tomz et al. (2007) and Subramanian & Wei (2007). Whenever a variable perfectly predicts an outcome, it is dropped. Moreover, observations are dropped also for countries which import from all other countries.

Table 2 reports the results, with different time periods lined up as different columns. All coefficients come with the expected signs whenever they are statistically significant. Comparing across time periods, we may note that distance becomes ever less important as a

[28] Liu (2007) tries to solve this problem by proxying unobserved variables.

[29] Rose (2004a) and Tomz et al. (2007) use total trade. Subramanian & Wei (2007) also use imports. However, although they use the same sources for their trade data include only 6,638 country pairs with strictly positive imports (for 2000).

Table 2: WTO membership and the extensive margin of trade

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Pre Kennedy		Kennedy-Tokyo		Tokyo-Uruquay		Post Uruguay	
Membership status:	formal	factual	formal	factual	formal	factual	formal	factual
Dependent variable:	$I_{(Mij>0)}$	$I_{(Mij>0)}$	$I_{(Mij>0)}$	$I_{(Mij>0)}$	$I_{(Mij>0)}$	$I_{(Mij>0)}$	$I_{(Mij>0)}$	$I_{(Mij>0)}$
WTO x IND	0.462***	0.00356	0.127***	0.0712***	0.00848	0.0580**	0.0440	0.0700
	(0.105)	(0.0524)	(0.0213)	(0.0254)	(0.0398)	(0.0292)	(0.0424)	(0.0438)
WTO x DEV	0.0192	-0.0212	0.0403***	0.0177	0.0131	0.0395***	0.0311**	0.0199
	(0.0475)	(0.0350)	(0.0138)	(0.0178)	(0.0108)	(0.0122)	(0.0155)	(0.0173)
RTA			0.504*	0.537*			0.480**	0.478**
			(0.302)	(0.318)			(0.214)	(0.213)
GSP	1.707	1.772	0.234***	0.263***	0.250***	0.241***	-0.0757	-0.0786
	(1.227)	(1.239)	(0.0309)	(0.0301)	(0.0599)	(0.0595)	(0.0810)	(0.0790)
Currency union	0.680***	0.674***	0.365***	0.367***	0.0999	0.101	0.495***	0.496***
	(0.169)	(0.168)	(0.0832)	(0.0845)	(0.0617)	(0.0617)	(0.186)	(0.187)
ln Distance	-0.389***	-0.391***	-0.171***	-0.176***	-0.128***	-0.127***	-0.0759***	-0.0760***
	(0.0145)	(0.0146)	(0.00698)	(0.00714)	(0.00589)	(0.00604)	(0.00400)	(0.00402)
Contiguity	0.122*	0.125*	-0.0159	-0.0197	-0.0163	-0.0196	-0.0317	-0.0314
	(0.0633)	(0.0641)	(0.0459)	(0.0475)	(0.0380)	(0.0386)	(0.0353)	(0.0352)
Common language	0.0563**	0.0586***	0.0273***	0.0300***	0.0459***	0.0450***	0.0285***	0.0284***
	(0.0220)	(0.0221)	(0.00936)	(0.00952)	(0.00680)	(0.00683)	(0.00475)	(0.00475)
Every colony	0.372***	0.378***	0.0805***	0.0812***	0.0279	0.0225		
	(0.0455)	(0.0479)	(0.0201)	(0.0209)	(0.0646)	(0.0671)		
Common colonizer	0.312***	0.318***	0.0896***	0.0921***	0.0601***	0.0587***	0.0217***	0.0215***
	(0.0185)	(0.0193)	(0.00686)	(0.00706)	(0.00610)	(0.00621)	(0.00501)	(0.00502)
Same country	-0.0635	-0.0686	0.00992	0.00853	0.0773***	0.0788***	0.0458***	0.0460***
	(0.0947)	(0.0939)	(0.0462)	(0.0478)	(0.0173)	(0.0164)	(0.0119)	(0.0117)
N	12147	12147	14131	14131	16453	16453	15212	15212
r2_p	0.560	0.559	0.491	0.490	0.423	0.423	0.440	0.440
obs. P			0.698	0.698	0.745	0.745	0.805	0.805
pred. P			0.879	0.875	0.903	0.903	0.944	0.944
chi2	3699	3700	4008	4000	4170	4155	3697	3687
Chi2 (IND = DEV)	17.75***	0.22	17.1***	5.806**	0.01	0.44	0.1	1.44

All coefficients are to be read as marginal effects (evaluated at sample averages). Robust standard errors in parentheses (corrected for clustering at country-pair level), *** p<0.01, ** p<0.05, * p<0.1. All specifications include importer and exporter effects and a constant (not shown). Pre-Kennedy refers to averages for period 1948-1967; Kennedy.Tokyo to 1968-1978; Tokyo-Uruquay to 1979-1994; Post Uruguay to 1995-2004. Number of observation reflects observations that are not perfectly predicted by covariates. Chi2 (IND=DEV) tests equality of coefficients WTO x IND and WTO x DEV.

determinant of the extensive margin of world trade. This may reflect the fact that countries' trading relationships are becoming more and more far-reaching geographically. This is an interesting result that has not been noted before. It contrasts with the "distance-puzzle" found in intensive-margin-models surveyed in Disdier & Head (2008), by which distance plays an increasingly important role in restricting trade.[30] Also, the trade-creating effect of colonial variables is weakening over time.

Industrial versus developing importers. Odd-numbered columns report the marginal effect on bilateral trade of the importer and the exporter both being *formal* members of the WTO. In the pre-Kennedy-round period, formal membership has had a fairly strong positive

[30] On the extensive margin, see Felbermayr & Kohler (2006).

impact on the likelihood of an industrialized importer having an active trade relationship with other WTO members. Membership raises those odds by about 46 percentage points. This is a large effect compared to other determinants of the probability of positive trade, such as the existence of a common language, which increases this likelihood by a mere 6 percent. On the other hand, granting easier market access to industrial countries through the generalized system of preferences (GSP) does not affect the likelihood of trade. The same is true if the importer WTO member is a developing country.

In the aftermath of the Kennedy-round, developing importers also see some effect of GATT membership (formal or informal) on the extensive margin, but the impact is much smaller than for industrialized countries (equality of coefficients is rejected at the 1 percent level). However, in the period between the conclusion of the Tokyo and the Uruguay rounds positive effects appear restricted to non-formal membership. Formal membership alone has no significant effect either for industrial or developing countries. After 1995, the extensive margin effect of formal membership seems to have become operative again for developing countries, although its quantitative importance is rather subdued. Overall, Table 2 tells that to some extent the conclusions drawn by Subramanian & Wei (2007) for the intensive margin carry over to the extensive margin. However, there is a reversal over time: Subsequent to the Uruguay-round, developing countries tend to benefit more than industrial countries.

Formal versus factual membership. Even-numbered columns in Table 2 define WTO membership on *factual* rather than formal grounds. Our results suggest that in the pre-Kennedy-round era, where the number of informal members was almost as large as the number of formal members, factual membership did not enhance the likelihood of trade. The same holds true in the interval between the Kennedy- and the Tokyo-round, where factual membership does not add anything to the likelihood of an active trading relationship. Things are different, however, for the third time phase between the Tokyo- and the Uruguay-round, where *factual* membership turns out to increase the odds of trade, while *formal* membership does not. Finally, in recent years, with informal members massively converting into formal ones, factual membership does not appear to affect the likelihood of trade, while formal membership does. Hence, we do not find that a focus on factual (as opposed to formal) membership makes the advantage of being associated to the GATT/WTO any more visible in an econometric analysis of the extensive margin. This is an important piece of new information relative to Tomz et al. (2007) who find this type of effect to obtain at the intensive margin. Our results are quite plausible: What would be the benefits of formal membership beyond informal one when the latter were at least as effective in activating trade relationships?

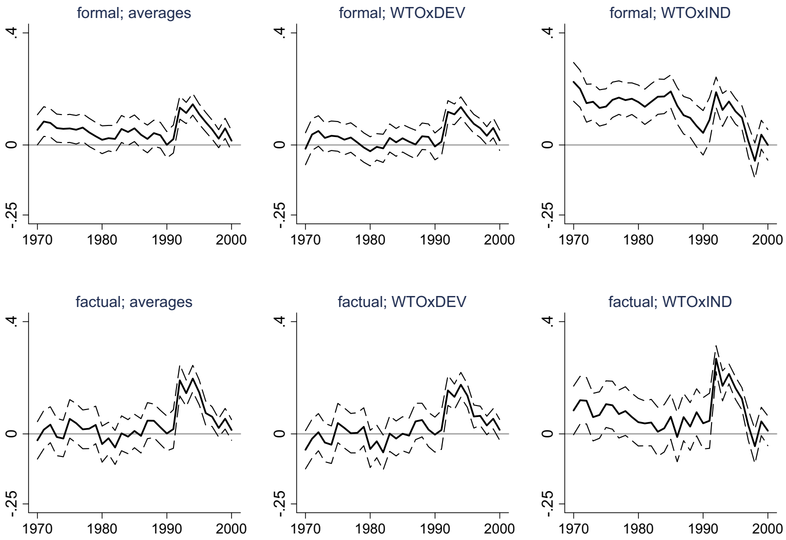

Figure 3: The extensive margin year-by-year (dashed lines indicate 5% confidence bands).

Robustness checks. Figure 1 plots WTO coefficients obtained by running regressions of the type discussed in Table 3 as yearly cross-sections rather than taking averages over the different phases of the GATT/WTO history. This procedure drastically increases the number of zeros in the data, but presumably comes at the cost of substantial measurement error.[31] The top row of panels depicts the effects of formal membership, while the bottom row depicts those of factual membership. Treating industrialized and developing countries alike, formal membership appears as a more robust positive influence on the probability of trade than factual membership. Both yield large positive results for some years after 1991. This may reflect the increased number of potential trading relationships as new countries have emerged due to the break-up of the URSS. Turning to developing countries, in earlier times formal WTO membership was worth much less for developing countries than for industrialized ones. However, this pattern has recently reversed. Moreover, we do not find any evidence that factual membership is more robustly associated with increased trading probabilities. If anything, the opposite holds true. Thus, year-by-year regressions tend to confirm the findings presented in Table 3.

We should, however, bear in mind that the amount of trade creation at the extensive margin is over the entire evolution of world trade after formation of the GATT has been relatively small; see Felbermayr & Kohler (2006) and Helpman et al. (2008). To eliminate the "pseudo-extensive margin" due to formation of new countries, Felbermayr & Kohler (2006) use a sample of countries that have been in existence ever since 1965. They find that

[31]To limit measurement problems, we focus on the post-1970 era, where idiosyncratic switches between zeros, positive trade values, and missings are far less frequent than in the years before.

15 percent of total world trade in 2004 is due to bilateral relationships that have not existed in 1965. Hence, while WTO membership may indeed increase the odds that two members engage in trade, the contribution of the extensive margin on total world trade growth cannot be possibly very large. In the next section we therefore study the intensive margin of world trade, but allowing for trade to be zero, as in the corner solutions version of the gravity model presented in section 3 above. Using a PPML estimation approach allows us to see to what extent the extensive margin changes the total pro-trade effect of WTO membership.

6.2 Estimating the full non-linear model

We have argued in section 4 that the potential of zero trade as an equilibrium outcome gives rise to a non-linear gravity model of bilateral trade. Following Santos & Tenreyro (2006), we now proceed to estimating this model using a Poisson pseudo maximum-likelihood (PPML) procedure.

Average effects: OLS versus PPML: Table 3 sets estimates from the standard log-log OLS procedure against Poisson pseudo maximum-likelihood (PPML) estimates. Columns labeled (A) present conventional results where the dependent variable is the log of bilateral imports, $\ln M_{ij}$. Columns labeled (B) use PPML on the same sample as in (A), while columns (C) extend the sample to zero-trade observations, which drop out in the OLS approach as the log of zero is not defined.

The OLS results look familiar. The elasticity of distance is fairly large and doubles in absolute terms over time. Cultural proximity (colonial ties, common language) boost trade considerably. *Formal* GATT/WTO membership increases bilateral trade in all time-spans considered, except for the time between the Kennedy- and the Tokyo-round of the GATT.

Running PPML on the same sample (i.e., excluding zero-trade observations), the distance coefficient is much smaller, the dummy for colonial ties is less frequently significant, and the regional trade agreement dummy comes with comparable size. As to WTO/GATT membership, the results are somewhat bleaker than those from OLS estimation. Indeed, during the first two time-spans considered, GATT membership seems associated with *lower* bilateral trade. This effect is quantitatively substantial and roughly constant over the two periods. In the Tokyo- to Uruguay-round era, WTO membership simply does not seem to matter. It is only in post-Uruguay-round times that abiding by rules of the WTO appears to boost bilateral trade. The effect is statistically significant and economically important. On average, the volume of trade is twice as large for WTO members than for outsiders.

Table 3: Trade creation and the extensive margin: OLS versus PPML

	(A)	(B)	(C)	(A)	(B)	(C)	(A)	(B)	(C)	(A)	(B)	(C)
		Pre Kennedy			Kennedy-Tokyo			Tokyo-Uruguay			Post Uruguay	
Dep.var.	ln Mij OLS	Mij>0 PPML	Mij PPML	ln Mij OLS	Mij>0 PPML	Mij PPML	ln Mij OLS	Mij>0 PPML	Mij PPML	ln Mij OLS	Mij>0 PPML	Mij PPML
WTO	0.235**	-0.690***	-0.735***	-0.193**	-0.853***	-0.886***	0.200**	-0.126	-0.177	0.836***	1.118***	1.113***
	(0.101)	(0.167)	(0.159)	(0.0966)	(0.183)	(0.179)	(0.0909)	(0.209)	(0.210)	(0.188)	(0.303)	(0.302)
RTA	-1.759**	0.0631	0.0118	0.982**	0.248	0.239	0.914***	0.146	0.143	0.389	0.562***	0.555***
	(0.883)	(0.349)	(0.350)	(0.426)	(0.161)	(0.161)	(0.224)	(0.123)	(0.123)	(0.295)	(0.159)	(0.159)
GSP	3.242***	-8.859***	-8.341***	1.678***	0.345***	0.383***	1.335***	0.247***	0.260***	1.305***	0.378***	0.396***
	(1.163)	(2.645)	(2.637)	(0.0736)	(0.0777)	(0.0781)	(0.0460)	(0.0553)	(0.0554)	(0.0793)	(0.0995)	(0.100)
Currency union	1.387***	1.896***	1.979***	1.327***	0.812***	0.861***	1.215***	-0.469	-0.452	2.720***	-1.860**	-1.857**
	(0.272)	(0.282)	(0.282)	(0.267)	(0.200)	(0.200)	(0.328)	(0.353)	(0.357)	(0.595)	(0.771)	(0.779)
ln Distance	-0.886***	-0.594***	-0.593***	-1.375***	-0.708***	-0.716***	-1.566***	-0.693***	-0.694***	-1.598***	-0.681***	-0.681***
	(0.0322)	(0.0462)	(0.0466)	(0.0322)	(0.0360)	(0.0364)	(0.0268)	(0.0330)	(0.0330)	(0.0247)	(0.0288)	(0.0289)
Contiguity	0.296**	0.314***	0.306***	0.328**	0.247**	0.230**	0.506***	0.410***	0.409***	0.434***	0.422***	0.425***
	(0.128)	(0.117)	(0.118)	(0.140)	(0.101)	(0.101)	(0.112)	(0.0853)	(0.0856)	(0.120)	(0.0825)	(0.0833)
Common language	0.472***	0.537***	0.579***	0.501***	0.424***	0.441***	0.493***	0.435***	0.443***	0.562***	0.285***	0.290***
	(0.0667)	(0.0961)	(0.0968)	(0.0677)	(0.0838)	(0.0846)	(0.0589)	(0.0819)	(0.0818)	(0.0541)	(0.0776)	(0.0777)
Ever colony	1.415***	0.802***	0.796***	1.514***	0.405***	0.414***	1.213***	0.133	0.132	1.072***	0.0812	0.0783
	(0.119)	(0.140)	(0.142)	(0.113)	(0.105)	(0.107)	(0.103)	(0.0985)	(0.0988)	(0.107)	(0.102)	(0.103)
Common colonizer	0.848***	0.0234	-0.0443	0.812***	-0.0886	-0.0715	0.965***	0.153	0.141	0.874***	-0.0232	-0.0194
	(0.0957)	(0.191)	(0.185)	(0.0943)	(0.148)	(0.149)	(0.0771)	(0.174)	(0.173)	(0.0695)	(0.197)	(0.197)
Same country	0.691***	-0.607**	-0.555**	1.032***	0.528***	0.414**	0.669***	0.519**	0.543**	0.545***	0.445***	0.451***
	(0.206)	(0.261)	(0.268)	(0.208)	(0.152)	(0.204)	(0.159)	(0.251)	(0.252)	(0.164)	(0.142)	(0.141)
Current colony	-0.668	1.896***	1.979***	0.228	1.229***	1.346***	-0.160	1.900***	1.904***	-0.212	1.774**	1.793**
	(0.914)	(0.282)	(0.282)	(0.774)	(0.326)	(0.347)	(1.057)	(0.501)	(0.503)	(1.309)	(0.701)	(0.706)
Observations	6301	6301	13155	10101	10101	14364	16186	16186	20388	19542	19542	22516
R-squared	0.735			0.751			0.762			0.779	.	.
F / chi2	85.09	29114	29114	150.1	97361	99282	198.9	61488	63775	249.1	73775	90285

Robust standard errors in parentheses (corrected for clustering at country-pair level), *** p<0.01, ** p<0.05, * p<0.1. All specifications include importer and exporter effects and a constant (not shown). Pre-Kennedy refers to averages for period 1948-1967; Kennedy.Tokyo to 1968-1978; Tokyo-Uruquay to 1979-1994; Post Uruguay to 1995-2004. Number of observation reflects observations that are not perfectly predicted by covariates.

The role of the extensive margin: Columns labeled (C) include zero-trade observations, thus allowing for WTO membership to enhance the likelihood of a trading relationship to exist at all, in addition to increasing the magnitude of trade. Including zero trade country pairs more than doubles the number of observations in earlier years. In more recent periods, due to the fact that increased utilization of potential trade relationships has reduced the number of zeros, the number of observations still increases by 15 percent.[32] The OLS results in columns (A) suffer from two problems. First, the log-log model is consistent only under very strong assumptions on the specification of the error term. And secondly, the results may be biased as the extensive margin is ignored. Columns (B) remedy the first problem, columns (C) help with both issues. Since (B) and (C) look fairly similar, it is save to conclude that the extensive margin *per se* does not play an important role in explaining the difference between OLS and PPML. For the post-Uruguay period, WTO membership increases bilateral trade by about 205 (i.e., $e^{1.118} - 1$) percent, whether zeros are included or not. Note that this magnitude is comparable to the findings in Subramanian & Wei (2007). As in their analysis, the comprehensive inclusion of country-specific fixed effects guarantees that multilateral trade resistance (e.g., relative to all countries in the world) is duly taken into account. Given the

[32]Without taking averages, the prevalence of zeros would be much stronger.

fact that most countries today are WTO members, it is tempting to interpret our WTO coefficient as the cost of non-membership.

Our results are in contrast to the panel estimates reported by Liu (2007), which seem to suggest that including zeros increases the trade-creating potential of WTO membership by about 30 percent. The difference is potentially due to three features of our approach. First, we include importer- and exporter-country fixed effects, in order to control for multilateral trade resistance as a determinant of bilateral trade. This turns out to be crucial for the results obtained. Secondly, for reasons outlined above we average the data over several years prior to estimation, while Liu uses yearly observations. To see what this implies, figure 4 reports results from yearly regressions and compares estimates on the a sample restricted to non-zero trade with those from unrestricted samples for the years 1970 to 2001. It turns out that averaging as such is not responsible for the weak WTO effect at the extensive margin, at least not in cross-sections. This leaves Liu's reliance on variance across time (within-variance) in a panel estimation framework as a second explanatory factor for a seemingly robust positive role of WTO membership for the extensive margin of trade. However, as pointed out above, a close inspection of our data leads us to question the reliability of time-series evidence on the occurrence of zero trade.[33]

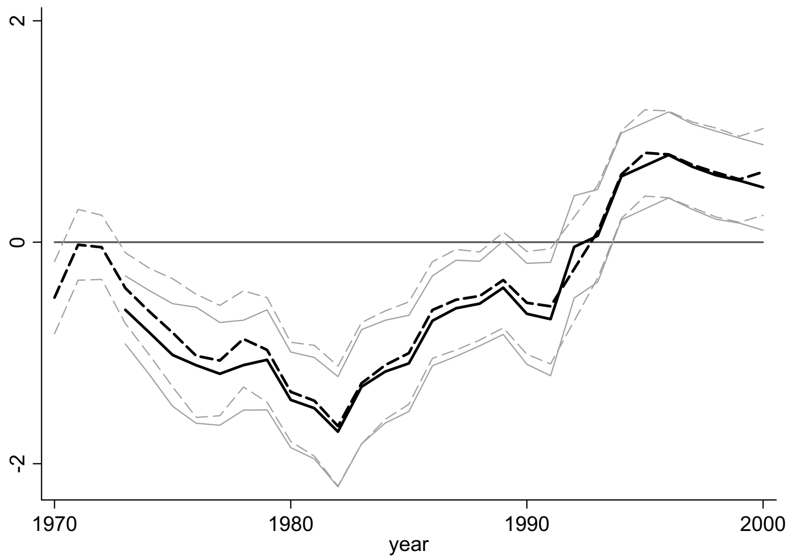

Figure 4: Restricted (solid lines) versus full (dashed lines) samples.

Industrial versus developing importers: Finally, Table 4 distinguishes between industrialized and developing WTO members, as well as between formal and factual participation

[33]His finding is hard to square with the low quantitative importance of new trade relationships in the 2004 total trade volume, see Felbermayr and Kohler (2006) and Helpman et al. (2008).

in the GATT/WTO. The sign of our coefficient estimates is consistently negative for the first three time spans considered and turn positive only in the post-Uruguay-round era. Concentrating on formal membership, we find that developing importers were not “penalized” by GATT membership in the pre-Kennedy-round era, while industrial importers were. In the Tokyo- to Uruguay-round period, this pattern reverses, while both types of countries seem to lose in the Kennedy- to Tokyo-round period. Turning to the period after the Uruguay agreement, we find strongly positive effects comparable in size to some of the estimates of Subramanian & Wei (2007). However, it turns out that developing countries benefit more strongly from WTO membership than industrialized countries. The difference in the coefficients is statistically significant and amounts to a differential trade creation of about 100 percent ($e^{1.343} - e^{1.035}$).

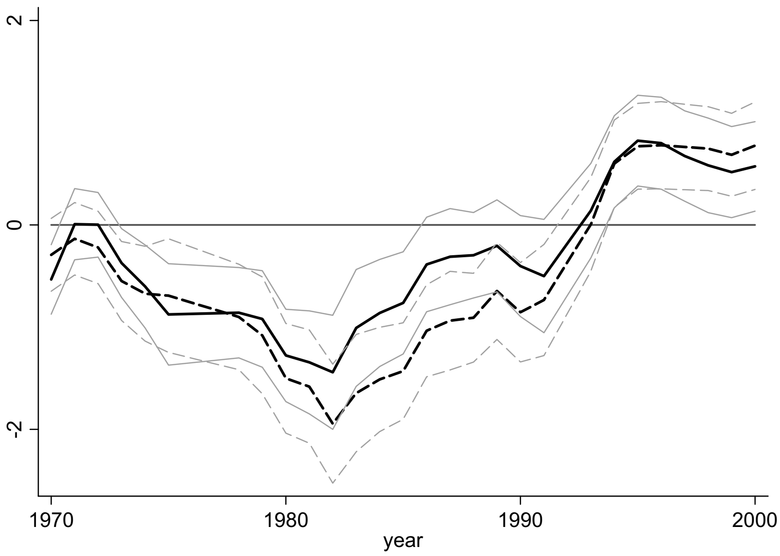

Figure 5: Industrialized (solid lines) versus developing (dashed lines) importers.

Overall, as with the Probit estimates on the extensive margin, there is no evidence that factual membership is more robustly associated with higher trade than formal one. In a similar vein, the comparison between industrial and developing countries does not depend on the definition of membership. Our results thus question the revisionist view proposed by Tomz et al. (2007).

Figure 5 provides a robustness check to the above findings, depicting the WTO coefficients from yearly regressions, where WTO membership is defined on formal grounds. It turns out that the coefficients for industrial and developing countries do not differ too much, except for the intermediary period, where industrialized countries have higher coefficients.

Table 4: Trade creation and the extensive margin: industrialized versus developing countries; formal versus factual membership

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Pre Kennedy		Kennedy-Tokyo		Tokyo-Uruquay		Post Uruguay	
Membership status:	formal	factual	formal	factual	formal	factual	formal	factual
Dependent variable:	Mij	Mij	Mij	Mij	Mij	Mij	Mij	Mij
WTO x IND	-1.060***	-0.974***	-0.906***	-1.034***	-0.0442	0.275	1.035***	1.017***
	(0.169)	(0.189)	(0.182)	(0.193)	(0.215)	(0.220)	(0.335)	(0.340)
WTO x DEV	-0.0515	-0.0144	-0.832***	-1.127***	-0.539**	-0.185	1.343***	1.285***
	(0.201)	(0.212)	(0.232)	(0.235)	(0.231)	(0.242)	(0.329)	(0.337)
RTA	0.00992	-0.0167	0.237	0.249	0.136	0.135	0.560***	0.561***
	(0.361)	(0.358)	(0.162)	(0.161)	(0.123)	(0.122)	(0.159)	(0.159)
GSP	-9.534***	-8.576***	0.368***	0.442***	0.328***	0.334***	0.362***	0.365***
	(2.718)	(2.649)	(0.0795)	(0.0775)	(0.0576)	(0.0575)	(0.116)	(0.116)
Currency union	1.875***	1.976***	0.854***	0.892***	-0.371	-0.380	-1.904**	-1.907**
	(0.282)	(0.283)	(0.200)	(0.198)	(0.349)	(0.335)	(0.777)	(0.774)
ln Distance	-0.622***	-0.609***	-0.716***	-0.713***	-0.694***	-0.693***	-0.682***	-0.681***
	(0.0471)	(0.0461)	(0.0364)	(0.0363)	(0.0331)	(0.0328)	(0.0288)	(0.0288)
Contiguity	0.314***	0.313***	0.232**	0.229**	0.407***	0.400***	0.425***	0.427***
	(0.119)	(0.117)	(0.101)	(0.100)	(0.0851)	(0.0841)	(0.0833)	(0.0834)
Common language	0.545***	0.572***	0.439***	0.424***	0.442***	0.432***	0.292***	0.292***
	(0.0963)	(0.0958)	(0.0848)	(0.0842)	(0.0820)	(0.0809)	(0.0764)	(0.0765)
Every colony	0.861***	0.843***	0.415***	0.418***	0.135	0.131	0.0773	0.0780
	(0.136)	(0.139)	(0.107)	(0.109)	(0.0989)	(0.0990)	(0.102)	(0.102)
Common colonizer	0.0452	-0.195	-0.0768	-0.00900	0.211	0.263	-0.0521	-0.0471
	(0.186)	(0.190)	(0.149)	(0.148)	(0.170)	(0.175)	(0.198)	(0.198)
Current colony	0.703*	1.311***	1.338***	1.396***	1.952***	1.978***	1.780**	1.715**
	(0.381)	(0.398)	(0.347)	(0.351)	(0.488)	(0.502)	(0.706)	(0.686)
Same country	-0.648**	-0.455*	0.406**	0.434**	0.578**	0.568**	0.442***	0.443***
	(0.272)	(0.259)	(0.206)	(0.201)	(0.247)	(0.248)	(0.141)	(0.141)
N	13155	13155	14364	14364	20388	20388	22516	22516
chi2	29114	29114	99396	99630	63529	64661	84388	85479
P-value IND = DEV								

Robust standard errors in parentheses (corrected for clustering at country-pair level), *** p<0.01, ** p<0.05, * p<0.1. All specifications include importer and exporter effects and a constant (not shown). Pre-Kennedy refers to averages for period 1948-1967; Kennedy.Tokyo to 1968-1978; Tokyo-Uruquay to 1979-1994; Post Uruguay to 1995-2004. Number of observation reflects observations that are not perfectly predicted by covariates.

7 Conclusion

World trade has evolved over the past five decades at both, the intensive and extensive margin. Existing studies exploring the role of GATT/WTO membership for trade have mostly been restricted to the intensive margin, ignoring country pairs with zero bilateral trade. The results are mixed, but trade-promoting influence of GATT/WTO membership remains in doubt. However, consistent estimation of the trade-promoting influence of GATT/WTO membership requires an empirical framework that allows for variation at the extensive margin that separates positive- from zero-trade country pairs. In this paper, we have developed what is probably the simplest possible gravity model that serves this purpose. We have estimated this model on a conventional data set of bilateral trade and the usual gravity controls, including in particular a set of dummy variables for trading arrangements, such as WTO membership.

Our estimation strategy duly responds to the issues raised in recent literature. Thus, we account for multilateral resistance, and we differentiate between country-groups (developing and industrial countries), as well as between formal membership as opposed to factual participation in the GATT/WTO. Moreover, responding to concerns about the reliability of trade data at the extensive margin, we abstain from panel estimation and rely on cross sectional evidence instead, averaging our data for the important phases of GATT/WTO history.

What is the conclusion that we may draw from our empirical exercise? First, we do find some evidence for WTO membership to raise the odds that countries trade with each other at all, but the effect is by no means robust across country groups and time. Nor does a broader definition of membership that includes de facto participation make much of a difference. The asymmetry, both across country groups and types of membership, changes erratically over sub-periods considered. A significantly positive Probit coefficient estimate emerges in less than half the number of specifications considered. The positive findings of WTO membership reported in Subramanian & Wei (2007) and Tomz et al. (2007) for the intensive margin thus do not carry over in any robust way to our Probit view on the extensive margin of world trade.

Things do not improve in any reassuring way, if we turn to a non-linear Poisson pseudo-maximum-likelihood (PPML) estimation of the non-linear model which looks at the intensive margin, but allows for zero-trade country pairs. Indeed, running PPML estimates on the entire sample of countries gives rise to an even bleaker picture for formal WTO membership than do conventional OLS estimates relying on positive trade only. It is only for the aftermath of the Uruguay-round that we obtain a significantly positive effect for formal membership. Grossly speaking, the same result obtains if we differentiate between types of membership and country groups.

The broad conclusion, then, is that the extensive margin does not prove a powerful line of defense for WTO membership as a trade-promoting force in a model which otherwise seems to work fine in terms of explanatory power, as well as the magnitude and significance of coefficient estimates.

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