

A RE-EXAMINATION OF THE IMPACT OF GATT/WTO ON TRADE

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Abstract

The empirical literature on the impact of membership in GATT/WTO on trade provides no conclusive results. The aim of this paper is to shed light on whether and to what extent GATT/WTO membership has increased world trade. We use traditional estimation techniques and recent developments in the econometric analysis of the gravity equation on a sample that covers 177 countries over the period 1960-2008. Our results show robust evidence that membership in GATT/WTO have had an economically significant effect on members' bilateral trade. Moreover, we find that the GATT/WTO effect operates through both trade margins but mainly through the intensive margin.

Key words: GATT; WTO; Trade; Gravity model; extensive margin; intensive margin

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1. Introduction

Since Rose (2004) seminal paper about the effect on international trade of multilateral trade agreements -the World Trade Organization (WTO) and its predecessor the General Agreement on Tariffs and Trade (GATT)-, several authors have investigated this issue with remarkably diverse results. Using the gravity model on a large panel dataset covering 178 countries over the period 1948-1999, Rose surprisingly finds that there is no evidence that GATT/WTO have increased world bilateral trade flows. This apparent inconsistency between conventional wisdom and Rose's results led this author to describe his results as an "interesting mystery".¹

Tomz et al. (2007) were the first that tried to solve this mystery. After updating Rose's dataset to include not only *de jure* but also *de facto* GATT/WTO membership, they conclude that the GATT/WTO substantially increased trade (about 70 per cent if both trading partners are GATT/WTO members).² In response to this article, Rose (2007) poses several concerns about the meaning, plausibility and robustness of their results and encourages for further research that addresses the question raised in these articles.

Some evident shortcomings of the above articles are related to the use of average bilateral trade data and the econometric specification estimated. In this sense, Subramanian and Wei (2007) focus on several asymmetries in the GATT/WTO system and on utilizing a properly specified empirical framework that controls for multilateral resistance terms. Using bilateral import flows from 1950 to 2000 (at five-year intervals)

¹ Since its inception in 1948, the GATT defined the rules of the world trade. Over near 50 years the GATT sponsored eight rounds of trade-policy negotiations that successfully reduced trade barriers (tariff rates firstly, and nontariff barriers later). The eighth round of talks (the Uruguay Round) led to the creation of the WTO in 1995. Moreover, over the years the GATT/WTO has also grown in the number of members from 23 (mainly developed) countries at the beginning to more than 150 nowadays. Therefore, given that multilateral trade liberalization is the aim of the GATT/WTO it seems reasonable to believe that the GATT and the WTO have had a major impact on world trade.

² In a related article (Goldstein et al., 2007) the same authors evaluate again the effect of the GATT/WTO, getting the same conclusion.

they find that the GATT/WTO promotes trade, strongly but unevenly. The unevenness is related to the asymmetries in the system. They find that the GATT/WTO boosts trade in industrialized countries, but not in developing countries; in less protected sectors, but not in agriculture and textile sectors; and for new WTO members, but not for old GATT members. Subramanian and Wei (2007), however, do not account neither for unobserved bilateral heterogeneity nor for differences in trade effects across preferential trade agreements (PTA).

Eicher and Henn (2011a) unify the Rose, Tomz et al. and Subramanian and Wei approaches with the aim of minimizing several potential omitted variable biases. Their framework controls comprehensively for three sources of omitted variable bias (multilateral resistance, unobserved bilateral heterogeneity and individual PTA trade effects) and they do not find evidence of positive WTO trade effects. Moreover, they show that multilateral resistance controls are suffice to negate WTO trade effects, concluding that all previous approaches produce the result that WTO membership does not generate statistically significant trade effects.

Dutt et al. (2011) document the effect of GATT/WTO membership on the (product-level) extensive and intensive margins of trade. Using 6-digit bilateral trade data, they find that the impact of WTO is concentrated almost exclusively on the extensive product margin of trade, i.e. trade in goods that were not previously traded. In particular, in their preferred specification, WTO membership increases the extensive margin of exports by 31 per cent whereas it has a negligible or even negative impact on the volume of already-traded goods.

Chang and Lee (2011) re-examine the GATT/WTO membership effect on bilateral trade flows using nonparametric methods. Their results suggest that membership in the GATT/WTO has large trade promoting effects that are robust to

several restricted matching criteria, alternative GATT/WTO indicators, non-random incidence of positive trade flows, inclusion of multilateral resistance terms and different matching methodologies.

Another strand of research highlights the sample selection bias in the traditional log-linear gravity formulation, derived from the fact that many country pairs exhibit zero trade flows. The papers discussed above use only the observations with positive trade and, therefore, these studies lose important information for assessing the impact of the GATT/WTO on trade. In line with this argument, Liu (2009) notes that by restricting the analysis to observations with positive trade flows, previous studies underestimate the effect of the GATT/WTO on trade. Using a fixed-effects Poisson quasi-maximum-likelihood estimator, Liu (2009) finds that the GATT/WTO membership boosts trade among members by 60 per cent (21 per cent through the extensive margin and 39 per cent through the intensive margin).

Finally, Herz and Wagner (2011a) also allow for zero trade flows using the fixed-effect Poisson maximum-likelihood estimator on a sample that covers the period 1953-2006 with annual data. They find that GATT/WTO promotes trade among members on average by 86 per cent, while trade with non-members is also fostered. However, an important caveat of this article is that it does not control for multilateral resistance terms (Anderson and van Wincoop, 2003).

In this paper we re-examine the impact of GATT/WTO on trade and trade margins using multiple parametric techniques, including several econometric approaches that deals with the presence of zeros in bilateral trade flows: The two-stage estimation procedure proposed by Helpman et al. (2008), the Poisson pseudo-maximum likelihood estimator suggested by Santos Silva and Tenreyro (2006 and 2010) and the fixed-effects Poisson maximum-likelihood estimator. Our sample covers 177 countries

over the period 1960-2008. To preview our results, we find robust evidence that GATT/WTO have had an economically significant effect on trade. Moreover, our results suggest that the GATT/WTO effect operates through both the extensive and the intensive margins, but it works mainly through a reduction in variable costs (intensive margin) rather than in fixed costs of trade.

The paper is structured as follows. Section 2 presents the methodology. Section 3 describes the data. Section 4 discusses the estimation results. Finally, section 5 concludes the paper.

2. Methodology

The gravity model is the key econometric technique used to analyse the determinants of bilateral trade flows and, in particular, to study the effects of economic integration agreements on bilateral trade flows.³ The standard gravity model of trade relates bilateral trade flows to GDPs, distance and other factors that affect trade barriers.⁴

Our benchmark specification is the following augmented gravity equation:

$$\begin{aligned} \ln X_{ijt} = & \beta_0 + \beta_1 \ln Y_{it} + \beta_2 \ln Y_{jt} + \beta_3 \ln D_{ij} + \beta_4 Cont_{ij} + \beta_5 Island_{ij} \\ & + \beta_6 Landl_{ij} + \beta_7 Lang_{ij} + \beta_8 Colony_{ij} + \beta_9 ComCountry_{ij} \\ & + \beta_{10} Creligion_{ij} + \beta_{11} CU_{ijt} + \beta_{12} PTApIur_{ijt} + \beta_{13} PTAbil_{ijt} \\ & + \beta_{14} UPR_{ijt} + \beta_{15} GATT / WTO_{ijt} + u_{ijt} \end{aligned} \quad (1)$$

³ The gravity model has been regularly used to estimate the impact of preferential trade agreements (see, for example, Baier and Bergstrand, 2007; Baier et al., 2007; Carrère, 2006; Gil et al., 2008a or Lee et al., 2008), currency unions (Gil et al., 2008b; Glick and Rose, 2002; Micco et al., 2003 or Rose, 2000), unilateral (nonreciprocal) preference regimes (Gil et al., 2011; Goldstein et al., 2007; Herz and Wagner, 2011b; Matoo et al., 2002; Rose, 2004; Subramanian and Wei, 2007) or, as in this paper, GATT/WTO membership (see the references cited in the introductory material).

⁴ The gravity equation is backed up by sound theory since the end of the 1970's. See, among others, Anderson (1979), Anderson and van Wincoop (2003), Bergstrand (1985 and 1989), Deardoff (1998), Eaton and Kortum (2002), and Evenett and Keller (2002).

where i and j denote trading partners, t is time, and the variables are defined as follows: X_{ijt} are the bilateral export flows from i to j in year t ,⁵ Y denotes Gross Domestic Product, D denotes the distance between i and j , $Cont$ is a dummy variable equal to one when i and j share a land border, $Island$ is the number of island nations in the pair (0, 1, or 2), $Landl$ is the number of landlocked areas in the country-pair (0, 1, or 2), $Lang$ is a dummy variable which is unity if i and j have a common language, $Colony$ is a binary variable which is unity if i ever colonized j or vice versa, $ComCountry$ is a binary variable which is unity if i and j were part of a same country in the past, $Creligion$ is an index of common religion⁶, CU is a binary variable which is unity if i and j use the same currency in year t , $PTAPlur$ ($PTABil$) is a binary variable which is unity if i and j belong to the same plurilateral (bilateral) preferential trade agreement, UPR is a binary variable which is unity if i is a beneficiary of an Unilateral Preference Regime and j is the corresponding preference-giving country, and $GATT/WTO$ is a binary variable which is unity if i and j participate in GATT/WTO, and u_{ijt} is the standard classical error term.

Equation (1) ignores the theoretical foundations for the gravity equation that have been developed since Anderson (1979) and, therefore, is likely mis-specified. In particular, Anderson and van Wincoop (2003) illustrate the omitted variables bias introduced by ignoring multilateral resistance (price) terms in gravity equations.⁷ As these authors emphasize, gravity model theory implies that one must take into account

⁵ A number of studies treat the average of two-way bilateral trade as the dependent variable (see, for example, Glick and Rose, 2002; Rose 2000 and 2004 or Tomz et al., 2007). Baldwin and Taglioni (2006) called this procedure as the *silver medal* mistake. All theories that underlie a gravity-like specification yield predictions on unidirectional bilateral trade rather than two-way bilateral trade. In this paper, we use unidirectional trade data and, therefore, our specification is more closely grounded in theory.

⁶ The index is defined as: $(\% \text{ Protestants in country } i * \% \text{ Protestants in country } j) + (\% \text{ Catholics in country } i * \% \text{ Catholics in country } j) + (\% \text{ Muslims in Country } i * \% \text{ Muslims in country } j)$.

⁷ Multilateral resistance captures the notion that trade decisions are based on relative, rather than absolute, prices.

not only the trade resistance between any two countries (the bilateral resistance, which is a function of distance, language, contiguity, etc.), but also the fact that different countries have different multilateral resistance to trade. In a cross-section framework, the usual solution to the presence of such multilateral resistance is to include country fixed effects (CFE) for both the exporter and the importer countries when estimating gravity equations. However, in a panel setting, separate country fixed effects should be included for each year as multilateral resistance may change over time (Baldwin and Taglioni, 2006). The specialised literature refers to these estimates as country-year fixed effects (CYFE).

More recently, Helpman et al. (2008) (henceforth HMR) extend the gravity model developed by Anderson and van Wincoop (2003) by adding controls for the presence of zero bilateral trade flows and for non-observable firm heterogeneity. Moreover, they also derive a two-stages estimation procedure to obtain the effects of trade barriers and trade policies on the intensive and the extensive margins of trade. In the first stage they estimate a probit equation that specifies the probability that country i exports to j conditional on the observable variables and uses it to estimate effects on the extensive margin. In the second stage, predicted components of this equation are used to estimate the gravity equation that allows them to obtain effects on the intensive margin. This procedure simultaneously corrects for two types of potential biases: a sample selection bias and a bias caused by firm heterogeneity.

More formally, in a first stage they estimate a probit equation of the type:

$$\text{Pr ob}(T_{ij} = 1 / \text{observed variables}) = \Phi(\chi_i, \lambda_j, X_{ij}, Z_{ij}, \varepsilon_{ij}) \quad (2)$$

where T_{ij} is an indicator variable equal to 1 when country i exports to j and zero when it does not, Φ is the cumulative distribution function of the standard normal distribution, χ_i and λ_j are exporter and importer fixed effects, X_{ij} are variables which affect both the

probability and the volume of trade, and Z_{ij} represents variables that are used for the exclusion restriction, that is, those that affect the probability of observing a positive volume of trade but do not impact the volume of trade if this were to be positive. Using the probit regression, they construct two variables that are included as regressors in the second stage estimation. One is the inverse of Mills ratio and the other is an expression that controls for firm size heterogeneity. In particular, the second stage consists in the estimation for a given year of the following non-linear equation for all country-pairs with positive trade flows:

$$\ln trade_{ij} = \beta_0 + \lambda_j + \chi_i - \gamma X_{ij} + \theta \hat{\eta}_{ij}^* + \ln \left\{ \exp \left[\delta (\hat{z}_{ij}^* + \hat{\eta}_{ij}^*) \right] - 1 \right\} + \varepsilon_{ij} \quad (3)$$

where $\hat{\eta}_{ij}^*$ is the inverse Mills ratio and $\hat{z}_{ij}^* = \Phi^{-1}(\hat{\rho}_{ij})$ in which $\hat{\rho}_{ij}$ are the estimates from the probit equation.⁸

Finally, Santos Silva and Tenreyro (2006 and 2010) focus on econometric problems resulting from heteroscedastic residuals and the prevalence of zero bilateral trade flows. These authors argue that both OLS as well as HMR two-stage estimators are biased in the likely presence of heteroskedasticity in trade data. Therefore, they propose a non-linear Poisson estimator to estimate the gravity equation which, in addition, accounts for the presence of zeros in bilateral trade flows.

3. Data

The trade data for the regressand (export flows from country i to country j) come from the “Direction of Trade” (DoT) dataset built up by the International Monetary Fund (IMF). The data comprise bilateral merchandise trade between 177 countries and territories (see Appendix B) for 13 years of the period 1960-2008 at four-year intervals

⁸ Since equation (3) is non-linear in δ , following HMR (2008) we estimate it using maximum likelihood.

(1960, 1964,...,2008).⁹ The DoT dataset provides FOB exports in US dollars. These series are converted into constant terms using the American GDP deflator taken from the Bureau of Economic Analysis (US Department of Commerce).

The independent variables come from different sources. GDP data in constant US dollars are taken from the World Development Indicators (World Bank). For location of countries (geographical coordinates), used to calculate Great Circle Distances, and the construction of the dummy variables for physically contiguous neighbours, island and landlocked status, common language, colonial ties, common religion and common country background data are taken from the CIA's World Factbook. The sample includes 294 preferential trade agreements and currency unions.¹⁰ The indicators of currency unions are taken from Reinhart and Rogoff (2002), CIA's World Factbook and Masson and Pattillo (2005). The indicators of preferential trade agreements have been built using data from the World Trade Organization, the Preferential Trade Agreements Database (Faculty of Law at McGill University) and the website <http://ec.europa.eu>. Moreover, the sample includes 15 unilateral preference regimes (10 GSP programs plus AGOA, EBA, Cotonou Agreement, CBI and APTA). The list of countries beneficiaries of the standard GSP programs are taken from the United Nations Conference on Trade and Development (UNCTAD, 2001, 2005, 2006 and 2008). For previous years, we use data from UNCTAD kindly provided by Bernard Herz and Marco Wagner. Data on AGOA and EBA come from the corresponding

⁹ It is noteworthy that not all the areas considered are countries in the conventional sense of the word. We also include some dependencies, territories and overseas departments in the data.

¹⁰ The expression PTAs in this paper refers also to other agreements involving a higher degree of economic integration. In fact, most economic integration agreements considered in the sample are free trade agreements. The list of PTAs and CUs are available from the authors upon request

websites¹¹. The list of beneficiaries of the Caribbean Basin Initiative (CBI) and the Andean Trade Preference Act (ATPA) come from the Office of United States Trade Representative. The listing of beneficiaries of the Cotonou Agreement comes from its website¹² and Head, Mayer and Ries (2010). Finally, data on membership in GATT/WTO come from World Trade Organization.

4. Empirical results

Our benchmark specification to estimate the impact of GATT/WTO is Ordinary Least Squares (with a full set of year-specific intercepts added to correct for common shocks and trends). The results are reported in column 1 of Table 1. The gravity equation works well in two senses. First, the equation fits the data well explaining near two-thirds of the variation of bilateral exports flows. Second, the estimated coefficients are, on the whole, intuitive in sign and size and both economically and statistically significant. The only exception is the negative estimated coefficient for the index of common religion. As it is usual, we find that economically larger countries trade more and more distant countries trade less. Moreover, the results indicate that landlocked countries trade less, whereas sharing a common border, a common language, colonial ties or the existence of islands in the country pair increase trade. Finally, we find a positive a statistically significant coefficients for the variables that capture the impact of the different economic integration agreements either nonreciprocal (unilateral) or reciprocal (bilateral, plurilateral and multilateral). In particular, the variable of interest (*GATT/WTO*), presents an estimated coefficient that is positive (0.068) and statistically insignificant at conventional levels.

¹¹ See, http://www.agoa.gov/eligibility/country_eligibility.html for membership in AGOA and <http://ec.europa.eu/trade/wider-agenda/development/generalised-system-of-preferences/everything-but-arms> for EBA.

¹² <http://ec.europa.eu/trade/wider-agenda/development/economic-partnerships>

Column 2 of Table 1 contains the regression results adding country-year fixed effects (CYFE) for exporters and importers to the benchmark equation. When CYFE are added, we no longer include controls that do not vary at the country-year level. In almost all cases, the impact goes in the same direction than in column 1. The exceptions are the estimated coefficients of the variables for common religion (that now is positive and statistically significant at the 1 per cent level) and bilateral PTAs (that loses the statistical significance at conventional levels). In particular, the estimated coefficient of the variable *GATT/WTO* (0.751) is again positive but much larger than that found without controls for multilateral resistance terms.

Time-varying country dummies (CYFE) should completely eliminate the bias stemming from the omission of multilateral resistance terms. The problem with this estimation is that it is not able to deal with unobserved bilateral heterogeneity, which is extremely likely to be present in bilateral trade flows and so, there may be omitted variables at the country-pair level that affect bilateral trade. In other words, time-varying country dummies do not remove the bias stemming from the correlation between the determinants of bilateral trade that have been included and the determinants that are unobservable to the researcher. Recognizing this, Baldwin and Taglioni (2006) and Baier and Bergstrand (2007) argue in favour of using time-invariant pair dummies in addition to time-varying country dummies. From the econometric discussion above, this is our preferred estimation technique for the sample that includes only positive values of trade. Results including country-year fixed effects and country-pair fixed effects reinforce our finding (column 3). Once again, the variable of interest presents an estimated coefficient that is positive (0.311) and statistically significant at the 1 per cent

level. Given that $\exp(0.311)$ equals 1.365, that coefficient implies that GATT/WTO, on average, increase trade by 36.5 per cent.¹³

In column (4) we consider using random effects at the country-pair level instead of fixed effects.¹⁴ This approach has the advantage of allowing the estimation of time-invariant variables. As we can see, the assumption of random effects strengthens the GATT/WTO trade effect. However, as it was expected, the Hausman specification test rejects the null hypothesis of no correlation between the individual effects and the explanatory variables suggesting that fixed effects are appropriate.

Columns 1 to 4 report the results for three specifications that include catch-all dummies for currency unions, preferential trade agreements (bilateral and plurilateral) and nonreciprocal preferential regimes. Eicher and Henn (2011a and 2011b), in papers on the measurement of the effect of currency unions and WTO membership on trade, respectively, show the importance of splitting the catch-all PTA and CU dummies into the individual PTAs and CU arrangements. According to these authors, if individual PTAs and CUs do not generate identical trade benefits, as a large empirical literature has documented, estimating an average coefficient using catch-all PTA or CU dummies generates biased results.

In line with the above argument, it is important to estimate the gravity equation allowing for individual effects for the different currency unions and the different

¹³ Equation is in logs. So, the percentage equivalent for any dummy is $[\exp(\text{dummy coefficient})-1]*100$.

¹⁴ In a panel framework, whether the fixed effects model (FEM) or the random effects model (REM) is the econometrically more appropriate setup depends on the potential correlation of the individual effects with the explanatory variables. If individual effects are correlated with the regressors only the FEM is consistent. However, if there is no such correlation the REM is both consistent and efficient. From an econometric point of view, the choice between FEM and REM must be based on the Hausman Test. Econometric evaluations to test for fixed versus random effects show wide evidence of the rejection of the REM. The choice of the fixed effects rather than random effects can also be justified on conceptual grounds since it is reasonable to believe that the source of endogeneity bias in the gravity equation is unobserved time-invariant heterogeneity (Baier and Bergstrand, 2007).

reciprocal and nonreciprocal trade agreements. The estimated coefficients of these variables and the fixed effects are not reported in the table for ease of presentation.¹⁵ According to the results reported in column 5 (CYFE) and 6 (CYFE & CPFE), the estimated coefficients do not change in a significant way and, in particular, the estimated coefficient of the variable of interest remains nearly unaltered with respect to those reported in columns 2 and 3, respectively. Thus, there seems to be robust evidence that multilateral trade liberalization have had a major impact on world trade.

The problem of all the above estimations is that in those regressions we use the sample of countries with positive trade volumes between them. Disregarding countries that do not trade with each other may produce biased estimates. Therefore, now we turn to the analysis of the results accounting for the presence of zero trade flows. To address this issue, in Table 2 we present the results using three estimations techniques: the two stages estimation procedure suggested by HMR (2008), the PPML estimator recommended by Santos Silva and Tenreyro (2006 and 2010) and the fixed effects Poisson maximum likelihood (PML) estimator used by Herz and Wagner (2011a and b).

In the first stage of HMR we estimate a probit equation to obtain the estimated probability of exporting and the effects of various trade barriers and trade policies on the extensive margin of trade. The results are presented in column 1. At this point, it is worth noting that the estimation of equation (2) might be subject to the incidental parameter problem, introducing a bias in the coefficients of the rest of variables (X_{ij} and Z_{ij}). However, as pointed out by Fernández-Val (2009), this bias does not affect the estimated marginal effects and, therefore, the predicted values obtained for the dependent variable. In general, the estimated marginal effects show the expected sign.

¹⁵ Our sample includes near 300 individual bilateral and plurilateral PTAs and CUs. For bilateral PTAs we have estimated an average coefficient using a catch-all dummy. The inclusion of individual dummies for bilateral PTAs does not affect the results in any significant way.

In particular, the estimated marginal effect of the *GATT/WTO* dummy is positive (0.147) and statistically significant at the 1 per cent level, suggesting that GATT/WTO have a trade-promoting effect on the extensive margin of trade, that is, they have created trade between countries that did not have trade relations before.

Using the probit regression, as explained before, we construct two variables for correcting sample selection bias and firm heterogeneity. The results for the second stage, which provides estimates of the effects of trade barriers and trade policies on the intensive margin, can be seen in column 2 of Table 2. The variable *CReligion* has been excluded from the estimation for identification reasons.¹⁶ Both the non-linear coefficient δ and the linear coefficient for $\hat{\eta}_{ij}^{*}$ are precisely estimated. The estimated coefficients of the remaining variables are similar to those found using OLS with CYFE, being the estimated coefficient of the variable of interest for the intensive margin 0.590.

The results of the PPML estimator appear in column 3 of Table 2. The regression fits the data well and explains 88 per cent of the variation in bilateral trade linkages. With the exception of the estimated coefficient of variable *CReligion* that is not statistically significant, the estimated coefficients are in line with those found using OLS and HMR estimators. In particular, the evidence about the positive impact of membership in GATT/WTO on trade remains unaltered.

The third and last way that we use to deal with the problem of the presence of zeros in bilateral trade flows is to apply the Poisson maximum likelihood (PML)

¹⁶ In this set-up, parameter identification requires the existence of a variable that affects the probability of observing a non-zero flow between two countries but not the volume. Following HMR (2008) we have used the variable common religion for this purpose. It is worth noting that the estimated coefficient of this variable with PPML (columns 3 and 6 of Table 3) is not statistically significant.

estimator with CPFE (column 4). The estimated coefficient of the variable of interest is, once again, positive (0.315) and highly statistically significant.¹⁷

Finally, following Baier and Bergstrand (2007) we account for “phased-in” and lagged terms-of-trade effects and we also test for *strict exogeneity*. To this end, we first introduce lagged effects of the economic integration agreements on trade. Results are in columns 1 and 2 of Table 4. We find that GATT/WTO has statistically significant lagged effects on trade flows being the cumulative average treatment effect (in the Baier and Bergstrand (2007)’s terminology) equal to 0.38 with one lag and to 0.42 with two lags. The economic interpretation of an average treatment effect of 0.42 is that after 8 years (our data are at four-year intervals) the GATT/WTO membership increases the level of trade by 52 per cent.

In order to test for the *strict exogeneity* of GATT/WTO, we include one lead to three alternative specifications (without lags, with one lag and with two lags). In our panel context if there is *strict exogeneity*, $GATT/WTO_{ij,t+1}$ should be uncorrelated with the contemporaneous trade flows. The results (columns 3 to 5 of Table 3) confirm this. The estimated coefficient of the variable $GATT/WTO_{ij,t+1}$ is not statistically significant at conventional levels in any case.¹⁸

5. Conclusions

The literature measuring the effect of GATT/WTO has produced mixed results. This paper re-examines this issue using traditional estimation techniques and recent developments in the econometric analysis of the gravity equation that deals with the

¹⁷ We also experimented by using PML with country-pair random effects and by excluding zeros in the poisson estimations (PPML and PML). In all the cases the results remained almost unchanged.

¹⁸ We also experimented with adding one or two lags and one lead to the dummy variables only to the variable $GATT/WTO$ and the results for the variable of interest did not change in any significant way.

presence of zeros in bilateral trade flows. We find robust evidence that GATT/WTO membership has had an economically significant effect on bilateral trade. When we use only the observations with positive trade, according to our preferred specification, GATT/WTO have increased trade by 36 per cent. After including the zero trade observations, we find a similar impact with the Fixed-effects Poisson maximum-likelihood estimator and larger estimated effects with HMR and PPML estimators.

Moreover, using the HMR two-stage estimation procedure, we find that the GATT/WTO had played an important role in creating trade at both the intensive and extensive margins. In particular, our results suggest that GATT/WTO effect operates by reducing primarily the variable costs (intensive margin) rather than the fixed costs of trade. This result is consistent with the fact that the volume of trade (once trading) is determined by variable costs which, in the international trade context, typically depend on distance and tariff barriers.

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Table 1. OLS and fixed effects estimations. Sample period 1960-2008 at four-year intervals.

Variables	(1) OLS	(2) CYFE	(3) CYFE & CPFE	(4) CYFE & CPRE	(5) CYFE	(6) CYFE & CPFE
LnY _{it}	1.047 (0.006)***					
LnY _{jt}	0.852 (0.007)***					
Ln D _{ij}	-1.060 (0.020)***	-1.284 (0.021)***		-1.263 (0.016)***	-1.262 (0.021)***	
Cont _{ij}	0.726 (0.084)***	0.496 (0.080)***		0.765 (0.067)***	0.512 (0.080)***	
Lang _{ij}	0.565 (0.041)***	0.470 (0.036)***		0.456 (0.033)***	0.423 (0.036)***	
Colony _{ij}	1.339 (0.118)***	1.485 (0.099)***		1.886 (0.087)***	1.439 (0.096)***	
ComCount _{ij}	2.509 (0.102)***	2.315 (0.126)***		2.333 (0.130)***	2.378 (0.175)***	
Island _{ij}	0.743 (0.098)***	0.694 (0.076)***		0.652 (0.060)***	0.631 (0.077)***	
Land _{ij}	-0.461 (0.028)***	-0.639 (0.063)***		-0.633 (0.054)***	-0.620 (0.061)***	
CReligion _{ij}	-0.374 (0.053)***	0.244 (0.048)***		0.199 (0.045)***	0.256 (0.049)***	
CU _{ijt}	0.671 (0.123)***	0.513 (0.110)***	-0.192 (0.067)***	-0.047 (0.063)		
RTAPlur _{ijt}	1.032 (0.055)***	0.900 (0.052)***	0.235 (0.035)***	0.689 (0.031)***		
RTABil _{ijt}	0.642 (0.056)***	0.033 (0.063)	0.069 (0.052)	0.123 (0.051)**		
UPR _{ijt}	0.125 (0.041)***	0.303 (0.048)***	0.232 (0.036)***	0.321 (0.032)***		
GATT/WTO _{ijt}	0.068 (0.025)***	0.751 (0.053)***	0.311 (0.036)***	0.769 (0.030)***	0.774 (0.053)***	0.314 (0.037)***
Time dummies	Yes	No	No	No	No	No
CYFE	No	Yes	Yes	Yes	Yes	Yes
CPFE	No	No	Yes	Yes	No	Yes
No observat.	114,997	134,718	134,718	134,718	134,718	134,718
Adj-R ²	0.61	0.70	0.37	0.69	0.71	0.37

Notes: The regressand is the log of real bilateral exports. Robust standard errors (clustered by country-pairs) are in parentheses. * significant at 10%; ** significant at 5%; *** significant at 1%. CYFE indicates time-varying exporter and importer fixed effects. CPFE indicates country pair fixed effects. The regressions in column (5) and (6) include individual dummies for each individual currency union, regional trade agreement and unilateral preferential regime.

Table 2. Estimation results accounting for the presence of zero trade flows. Sample period 1960-2008 at four-year intervals.

Variables	HMR two-stage estimation with CYFE		PPML with CYFE	PML with CPFE
	(1) Probit coefficients	(2) Marginal effects	(3) ML	(4)
LnY _{it}				1.027 (0.088) ^{***}
LnY _{jt}				0.781 (0.061) ^{***}
Ln Dist _{ij}	-0.781 (0.012) ^{***}	-0.291 (0.004) ^{***}	-1.093 (0.026) ^{***}	-0.623 (0.039) ^{***}
Cont _{ij}	0.108 (0.066) ^{***}	0.039 (0.023) ^{***}	0.193 (0.078) ^{***}	0.318 (0.068) ^{***}
Lang _{ij}	0.436 (0.019) ^{***}	0.151 (0.006) ^{***}	0.409 (0.036) ^{***}	0.176 (0.063) ^{***}
Colony _{ij}	0.803 (0.095) ^{***}	0.237 (0.019) ^{***}	1.154 (0.097) ^{***}	1.087 (0.084) ^{***}
ComCount _{ij}	1.237 (0.190) ^{***}	0.302 (0.021) ^{***}	1.955 (0.129) ^{***}	1.306 (0.184) ^{***}
Island _{ij}	0.308 (0.033) ^{***}	0.107 (0.011) ^{***}	0.633 (0.074) ^{***}	0.585 (0.130) ^{***}
Land _{ij}	-0.324 (0.029) ^{***}	-0.123 (0.011) ^{***}	-0.586 (0.061) ^{***}	-1.311 (0.190) ^{***}
CReligion _{ij}	0.190 (0.026) ^{***}	0.071 (0.010) ^{***}		0.035 (0.107)
CU _{ijt}	0.629 (0.108) ^{***}	0.197 (0.026) ^{***}	0.490 (0.106) ^{***}	1.310 (0.216) ^{***}
RTAP _{lurijt}	0.412 (0.038) ^{***}	0.140 (0.011) ^{***}	0.731 (0.052) ^{***}	1.273 (0.108) ^{***}
RTAB _{ilijt}	-0.137 (0.096)	-0.052 (0.037)	-0.035 (0.063)	-0.273 (0.119) ^{**}
UPR _{ijt}	0.250 (0.036) ^{***}	0.089 (0.012) ^{***}	0.380 (0.047) ^{***}	0.988 (0.144) ^{***}
GATT _{ijt}	0.399 (0.022) ^{***}	0.147 (0.008) ^{***}	0.590 (0.051) ^{***}	0.758 (0.183) ^{***}
δ			0.464 (0.035) ^{***}	
θ			1.246 (0.041) ^{***}	
No observat.	241,669		134,718	246,437
Pseudo R ²	0.51			0.88

Notes: Robust standard errors (clustered by country-pairs) are in parentheses. * significant at 10%; ** significant at 5%; *** significant at 1%. CYFE indicates time-varying exporter and importer fixed effects. The regressions include separate dummies for individual currency unions, regional trade agreements and unilateral preference regimes.

Table 3. Panel gravity equations with bilateral and time-varying country fixed effects

Variables	(1)	(2)	(3)	(4)	(5)
GATT _{ijt}	0.098 (0.045)**	0.085 (0.048)*	0.298 (0.047)***	0.125 (0.055)**	0.098 (0.059)*
GATT _{ij,t-1}	0.285 (0.040)***	0.144 (0.048)***		0.244 (0.044)***	0.112 (0.051)**
GATT _{ij,t-2}		0.186 (0.041)***			0.199 (0.047)***
GATT _{ij,t+1}			0.008 (0.057)	-0.005 (0.058)	0.006 (0.060)
CU _{ijt}	-0.328 (0.082)***	-0.296 (0.084)***	0.113 (0.113)	-0.165 (0.121)	-0.206 (0.124)*
CU _{ij,t-1}	0.190 (0.095)**	-0.103 (0.126)		0.382 (0.116)***	0.007 (0.143)
CU _{ij,t-2}		0.439 (0.112)***			0.548 (0.128)***
CU _{ij,t+1}			-0.437 (0.099)***	-0.445 (0.100)***	-0.271 (0.100)***
RTAP _{lurijt}	0.119 (0.048)**	0.085 (0.049)*	-0.067 (0.056)	-0.104 (0.064)	0.118 (0.066)*
RTAP _{lurij,t-1}	0.114 (0.049)**	0.185 (0.061)***		0.003 (0.060)	-0.011 (0.075)
RTAP _{lurij,t-2}		-0.195 (0.060)***			-0.094 (0.076)
RTAP _{lurij,t+1}			0.267 (0.053)***	0.263 (0.055)***	0.236 (0.057)***
RTAB _{ilijt}	0.018 (0.051)	0.121 (0.054)**	0.072 (0.056)	0.050 (0.059)	0.117 (0.060)**
RTAB _{ilij,t-1}	0.080 (0.052)	0.001 (0.054)		-0.006 (0.077)	0.016 (0.093)
RTAB _{ilij,t-2}		0.168 (0.072)**			-0.142 (0.112)
RTAB _{ilij,t+1}			-0.036 (0.053)	0.047 (0.055)	0.088 (0.057)
UPR _{ij,t}	0.019 (0.053)	0.051 (0.055)	0.260 (0.058)***	-0.016 (0.068)	-0.017 (0.069)
UPR _{ij,t-1}	0.371 (0.056)***	0.128 (0.064)**		0.498 (0.069)***	0.149 (0.091)*
UPR _{ij,t-2}		0.448 (0.060)***			0.492 (0.073)***
UPR _{ij,t+1}			-0.008 (0.054)	0.047 (0.055)	0.110 (0.057)*
No observat.	130,727	125,959	116,827	112,836	108,068
Adj-R ²	0.83	0.83	0.82	0.82	0.82

Notes: The regressand is the log of real bilateral exports. Robust standard errors (clustered by country-pairs) are in parentheses. * significant at 10%; ** significant at 5%; *** significant at 1%. CYFE indicates time-varying exporter and importer fixed effects. CPFE indicates country pair fixed effects.