DOES THE GATT/WTO PROMOTE TRADE? AFTER ALL, ROSE WAS RIGHT

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Abstract

This paper re-examines the GATT/WTO's trade impact using recent econometric developments that allow us estimating structural gravity equations with the Poisson pseudo-maximum likelihood (PPML) estimator on a large dataset that requires calculating high-dimensional fixed effects. By doing so, we overcome computational limitations that are present in previous studies. In line with Rose's (2004) seminal work, we find that, unlike regional trade agreements and currency unions, the GATT/WTO accession has not generated positive trade effects. This result is robust across periods and country groups; when using data at five-year intervals or for consecutive years; and when taking into account the GATT/WTO accession dynamics.

Key words: GATT/WTO; Trade; Gravity model; PPML; High-dimensional fixed effects.

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1. Introduction

Over the last 70 years the GATT and its successor from 1995, the WTO, have sponsored nine rounds of trade-policy negotiations that have successfully reduced trade barriers and contributed to a more transparent and predictable environment for world trade. Until the early 2000s there was a broad consensus on the important role played by GATT/WTO in promoting international trade. However, in a seminal empirical contribution, Rose (2004) reported the striking finding that countries acceding or belonging to the GATT/WTO did not have significantly different trade patterns than non-members. This contradiction between the conventional view and Rose's results led this author to describe his finding as an "interesting mystery" that deserved further research. Ever since, a considerable number of studies have attempted to solve this puzzle by updating Rose's dataset, accounting for potential sources of omitted variables bias, using alternative econometric techniques, taking into account the margins of trade (extensive and intensive) or splitting the sample by groups of countries, periods and sectors.

A review of the literature reveals that there has been an intense debate on this issue over more than a decade.[1] The empirical work relies on different specifications of the gravity equation. While some papers confirm Rose's finding (Eicher and Henn, 2011 and Roy, 2011), and there exists a large heterogeneity in the results across group of countries and periods, most papers find that, as a whole, the GATT/WTO has had a trade promoting effect in line with the aforementioned consensus view. Tomz *et al*. (2007), Liu (2009), Chang and Lee (2011), Herz and Wagner (2011), Dutt *et al*. (2013), Cheong *et al*. (2014), Kohl and Trojanowska (2015), Kohl (2015), and Gil-Pareja *et al*. (2016) find evidence of such a trade-enhancing effect. Moreover, other papers find

[1] See Gil-Pareja *et al*. (2016) for a comprehensive review of the empirical literature on the effect of GATT/WTO on trade.

evidence of a positive effect but only for some groups of countries, sectors or periods (Subramanian and Wei, 2007; Felbermayr and Kohler, 2010; and Bista, 2015).[2]

In parallel with this literature, several authors have focused on seeking the proper econometric specification for the gravity equation. Glick and Rose (2002), Egger and Pfaffermayr (2003), Chen and Wall (2005) or Baier and Bergstrand (2007) illustrate the importance of including time-invariant country-pair fixed effects to control for unobservable bilateral heterogeneity and endogeneity.[3] Baier and Bergstrand (2007) and Baldwin and Taglioni (2007) suggest that the gravity equation should also include exporter-time and importer-time fixed effects to control for changes in multilateral resistance terms (Anderson and van Wincoop, 2003). Last but not least, Santos Silva and Tenreyro (2006 and 2010) propose to use the Poisson Pseudo-Maximum Likelihood (PPML) estimator to deal with econometric problems resulting from heteroskedastic residuals and the prevalence of zeros in bilateral trade flows.[4]

Despite the fact that the available empirical literature on the effect of GATT/WTO on trade has progressively improved the econometric specifications to account for potential sources of bias, computational issues have so far conditioned the choice of estimator. The large datasets used in the estimation of the GATT/WTO effect (requiring to compute three different types of high-dimensional fixed effects) and/or

[2] Subramanian and Wei (2007) conclude that the GATT/WTO promotes trade strongly, but unevenly. In particular, 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. Moreover, Felbermayr and Kohler (2010) document a positive effect on trade for developing country importers in the post-Uruguay Round era. Finally, Bista (2015) finds a positive impact but only on the extensive margin in trade between industrial and developing members.

[3] Since Baier and Bergstrand (2007) pointed out that trade agreements are not exogenous, the endogeneity issue has received a great deal of attention in the empirical gravity-equation literature. These authors proposed the inclusion of country-pair fixed effects to deal with this problem. However, it is worth noting that country-pair dummies do not completely eliminate the extent of endogeneity. Therefore, this paper will test for strict exogeneity in section 4.

[4] Some recent papers (see, for example, Dai *et al*., 2014; Bergstrand *et al*., 2015; Anderson and Yotov, 2016; Baier *et al*., 2016; Matoo *et al.*, 2017) show the importance of including also the internal trade in the estimation of the gravity equation of international trade. Despite the algorithm developed by Larch *et al.* (2017) allows for it without loss of generality, as these authors and most empirical applications do, we do not include within-country trade flows. The unavailability of the required data (in terms of both countries and years) precludes this possibility.

difficulties to achieve convergence have precluded accounting simultaneously for unobserved bilateral heterogeneity and endogeneity (with country-pair fixed effects), multilateral resistance terms (with exporter-time and importer-time fixed effects), heteroskedastic residuals and zero trade flows.[5] However, recently Larch *et al*. (2017) have provided an iterative PPML estimator that accounts for all above issues in large datasets.[6]

This paper uses the computational development brought about by Larch *et al*. (2017) to estimate for the first time, to the best of our knowledge, the GATT/WTO effect on trade using the PPML estimator with the aforementioned three types of (high-dimensional) fixed effects. We carry out the estimations employing a dataset that includes trade flows between more than 200 countries over the period 1948-2013. Therefore, we need to compute more than 50,000 fixed effects to obtain unbiased, and theory-consistent estimates.

Our findings suggest that once we comprehensively account for all sources of bias cited above, the (direct) positive GATT/WTO trade effect vanishes, which is in line with Rose (2004). Moreover, the absence of positive effects is robust to different specifications. That is, it holds across time periods (using alternative classification criteria), when we distinguish both between early and late joiners, and between industrial and developing countries (independently of the direction of trade flows). Furthermore, in the case of developing countries, we do not find evidence of a positive effect associated with the regime change related with the Uruguay Round either. The results also remain nearly unaltered when including lags in the regression and they are

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[5] In this literature, five papers account for both heteroskedastic residuals and zeros using Poisson estimators (Liu, 2009; Felbermayr and Kohler, 2010; Herz and Wagner, 2011; Bista, 2015 and Gil Pareja *et al*. 2016) but none of them simultaneously controls for unobserved bilateral heterogeneity and multilateral resistance terms.

[6] Larch *et al*. (2017) apply this iterative PPML estimator to re-assess the currency union effect on trade concluding that the euro effect is economically small and statistically insignificant.

also robust to the use of data for consecutive years instead of data at five-year intervals. Interestingly, in contrast to the results for the GATT/WTO, we find strong support for the positive effect of regional trade agreements and currency unions on export flows.

The rest of the paper is organized as follows. Section 2 presents the methodology. Section 3 describes the data. Section 4 presents and discusses the results. Finally, section 5 concludes.

2. Methodology

Since it was independently developed by Tinbergen (1962) and Pöyhönen (1963) more than five decades ago, the gravity model has become the main econometric approach for the *ex post* estimation of the "partial" (or direct) effects of different kinds of economic integration agreements on bilateral trade, including the GATT/WTO. This paper uses the estimation strategy recently proposed by Larch *et al.* (2017).[7] This approach allows us estimating the gravity equation using PPML on a large dataset requiring to compute three types of high-dimensional fixed effects (exporter-year, importer-year and country-pair) to avoid biased estimates and misleading inference.

Baltagi *et al*. (2003), Baier and Bergstrand (2007), Baldwin and Taglioni (2007) and Gil-Pareja *et al.* (2008a,b) motivated and included the three types of fixed effects in the estimation of log-linear gravity equations of international trade. This set of fixed effects deals with two sources of omitted variables bias. On the one hand, country-pair fixed effects control for the impact of any time-invariant determinant of bilateral trade (observed or not) correlated with the regressors.[8] On the other hand, Anderson and van Wincoop (2003), in their theoretical foundation of the gravity equation, highlight that

[7] Zylkin's *ppml_panel_sg* command implements this procedure in Stata

[8] The argument is that there may be unobserved country-pair characteristics that affect trade, and which are at the same time correlated with the economic integration agreements. Baier and Bergstrand (2007) address this issue with respect to free trade agreements suggesting the use of dyadic fixed effects to avoid this omitted variable bias.

bilateral trade flows depend not only on bilateral trade barriers between any two countries but also on trade barriers of each country with the rest of the trading partners (i.e., the multilateral resistance).[9] They show that omitting a variable that reflects each country's multilateral resistance to trade leads to biased estimates. In a panel data setting, the usual solution to this problem is to include country-year fixed effects for both importers and exporters. Eicher and Henn (2011), Roy (2011), Dutt *et al*. (2013), Cheong *et al.* (2014), and Gil-Pareja *et al.* (2016) have estimated the effect of GATT/WTO on trade using log-linear structural gravity equations that control simultaneously for both unobserved bilateral heterogeneity (with country-pair fixed effects) and multilateral resistance terms (with exporter-time and importer-time fixed effects).

The PPML estimator, initially proposed by Santos-Silva and Tenreyro (2006) to fit the gravity model of bilateral trade flows, has two interesting properties when compared to the traditional log-linear gravity regression. First, it avoids the statistical problems that arise from the existence of zero bilateral trade flows.[10] Second, it solves econometric problems that emerge in the presence of heteroskedastic residuals. It is worth pointing out that the existence of heteroskedasticity affects both the efficiency and the consistency of an estimator and, as Santos-Silva and Tenreyro (2006) emphasize, this is the more important rationale for using PPML.

It is important to notice that this paper is not the first to address either zero trade flows or both zeros and heteroscedastic residuals in the GATT/WTO empirical literature. On the one hand, several articles estimate the GATT/WTO effect on trade taking into account zeros without dealing with the problem of heroskedasticity. The two

[9] Anderson (1979) and Bergstrand (1985) offer early theoretical justification for the gravity model.

[10] Obviously, the gravity equation in its log-linear specification is not defined for zero trade flows. This problem results in a sample selection bias that can be particularly important in datasets with a large number of trade observations that are zero in levels.

earliest papers in this group look at the GATT/WTO issue in a peripheral way. The first one, Felbermayr and Kohler (2006), relies on the Tobit model to incorporate zero trade flows. The second paper (Helpman *et al*., 2008) accounts for non-observable firm heterogeneity in a framework that also considers an extensive country-level margin of trade, running a Heckman-type procedure for empirical estimation. This second approach is also used, as a robustness check, by Dutt *et al.* (2013) in their work on the effect of WTO on the extensive and intensive product margins of trade. It is worth pointing out that both methods hinge crucially on the assumption of homoskedasticity.[11] Other articles that focused particularly on the case study of GATT/WTO address the problem of zeros with alternative approaches that are also subject to criticism. Roy (2011) includes zero trade observations by adding a small positive constant to all import flows to allow for log-linearization of zero trade flows.[12] Analogously, Kohl and Trojanowska (2015), include zero trade flows, by recoding them from 0 to 1. Finally, Kolh (2015) incorporates zero trade flows using (zero-inflated) negative binomial maximum likelihood estimation, a method that has been criticized because it depends on the unit of measurement of the dependent variable (Head and Mayer, 2014, p. 174).

On the other hand, some articles both account for zeros and also allow for heteroskedastic residuals using a Poisson estimator. The first paper that estimates the GATT effect on trade dealing with both problems at once is Liu (2009). Felbermayr and Kohler (2010), Herz and Wagner (2011), Bista (2015) and Gil-Pareja *et al*. (2016) have subsequently pursued the Poisson approach. However, none of them include country-

[11] Tobit and Heckman-type procedures can deal with zero trade relationships but they are not robust to misspecification of the error term (Felbermayr and Kohler, 2010).
[12] Santos Silva and Tenreyro (2006) show that this approach leads to inconsistent parameter estimates.

pair fixed effects and country-year fixed effects in the gravity equation simultaneously due to convergence issues or because the large number of fixed effects precludes it.[13]

Hence, this paper contributes to this literature by estimating the following gravity equation using PPML:

$$X_{ijt} = \exp(\beta_1 RTA_{ijt} + \beta_2 CU_{ijt} + \beta_3 GATT/WTO_{ijt} + \chi_{it} + \lambda_{jt} + \eta_{ij}) + u_{ijt} \qquad (1)$$

where i denotes the exporter, j denotes the importer and t is time. The dependent variable is the value of bilateral export flows (in levels), and the set of independent variables includes binary dummy variables for common membership in regional trade agreements (*RTA*), currency unions (*CU*) and *GATT/WTO* (our variable of interest), as well as exporter-time fixed effects (χ_{it}), importer-time fixed effects (λ_{jt}) and country-pair fixed effects (η_{ij}).[14] Finally, u_{ijt} denotes the error term.

Furthermore, some robustness checks are carried out by examining the impact of GATT/WTO across periods and groups of countries by splitting the variable of interest in gravity equation (1) accordingly.

3. Data

This paper uses Glick and Rose (2016) dataset and extends it by including the GATT/WTO dummy variables.[15] The data comprise bilateral trade flows between more

[13] Larch *et al.* (2017) provides a list of papers on other areas of research that are unable to obtain estimates with a full set of fixed effects with PPML.
[14] It is worth noting that the reference category for the economic integration agreements dummy variables (*RTA*, *CU* and *GATT/WTO*) include both pairs of non-member countries and member-non-member pairs avoiding the concern about multicollinearity raised by Cheong *et al.* (2014).
[15] We gratefully acknowledge Andrew Rose for making his data public.

than 200 IMF country codes over the period 1948-2013 (with gaps).[16] The dependent variable (bilateral exports flows in US dollars) comes from *Direction of Trade* dataset assembled by the International Monetary Fund. Data on GDPs come from *World Development Indicators,* supplemented where necessary by *Penn World Table* Mark 7.1 and IMS's *International Financial Statistics*. The data for latitude and longitude, landlocked and island status, physically contiguous neighbors, language and colonizers have been obtained from CIA's *World Factbook.* Currency Union data rely on the IMF's *Schedule of Par Values* and issues of the IMF's *Annual Report on Exchange Rates Arrangements and Exchange Restrictions*, supplemented with information from the *Statesman's Yearbook.* Following Glick and Rose (2016), we use a transitive definition of currency union. That is, if dyads *x*-*y*, and *x*-*z* are in currency unions, then *y*-*z* is a currency union. Data on regional trade agreements are taken from the World Trade Organization's website. We also resort to this website to obtain the date of accession of each country to the multilateral trade system used to create the dummy variables for GATT/WTO membership.

4. Empirical results

As a benchmark, Table 1 presents the results from three estimators that have been widely employed in previous studies on the effect of GATT/WTO on trade, which do not simultaneously account for all sources of estimation bias discussed above.[17] The first one is the OLS estimator with time-varying exporter and importer fixed effects as

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[16] It is noteworthy that not all areas covered are countries in the conventional sense of the word. The dataset also includes some colonies (e.g. Gibraltar), territories (e.g. Guam) and overseas departments (e.g. Guadeloupe).

[17] We use data at five-year intervals as in Chen and Wall (2005); Baier and Bergstrand (2007); Subramanian and Wei (2007); Eicher and Henn (2011); Behar, Cirera-i-Crecillé (2013); and Kohl (2014). Alternatively, Dai *et al*. (2014); Bergstrand *et al.* (2015) and Gil-Pareja *et al*. (2016) use of data for every four years. The use of data at intervals addresses the concern raised by Chen and Wall (2005, p. 52): "Fixed-effects estimation is sometimes criticized when applied to data pooled over consecutive years on the grounds that dependent and independent variables cannot fully adjust in a single year's time."

well as time-invariant country-pair fixed effects. The second one is the (country pair) Fixed-Effect Poisson maximum-likelihood estimator. The third one is the PPML estimator with time varying, directional (source and destination) country-specific dummies.

The results for the log-linear version of the gravity equation with OLS appear in column 1 of Table 1. At first glance, the estimated coefficients for the three types of economic integration agreements (regional trade agreements, currency unions and the GATT/WTO) are positive and statistically significant at conventional levels. In particular, the point estimate for the *GATT/WTO* variable is 0.162 with a standard error of 0.031, implying that GATT/WTO entry expands trade by 17.6 percent [exp(0.162)-1=0.176]. However, it is worth pointing out that this estimator does not tackle the issues related to heteroskedasticity and zeros.

Column 2 provides the results using the Poisson estimator with country-pair fixed effects, which accounts for heteroskedastic residuals, zeros and unobserved bilateral heterogeneity but not for multilateral resistance terms since it does not include exporter-time and importer-time fixed effects.[18] The estimated coefficients for the GDPs are in line with those reported in previous studies. As in column 1, the point estimates of the three economic integration agreements are positive and highly statistically significant. In this case, the point estimate for the GATT/WTO raises to 0.224 with a standard error of 0.052.

Finally, column 3 of Table 1 presents the results when we control for heteroskedastic residuals, zeros and the multilateral resistance, but neither for endogeneity nor for unobserved bilateral heterogeneity with country-pair fixed effects.

[18] In this specification we include the logarithm of the GDP of the exporter and the importer. The presence of exporter-time and importer-time fixed effects in all the other specifications in this paper captures any exporter specific and importer specific time-variant variable (such as GDPs) as well as all other time-varying country-specific unobservables affecting trade, including the theoretical multilateral resistance terms.

However, in this specification, we include bilateral time invariant trade supporting or impeding measures. In particular, we include the logarithm of bilateral distance (*Dist*), as well as dummy variables for adjacency (*Cont*), the use of a common language (*Lang*), the existence of colonial ties (*Colony*), being a common country in the past (*ComCount*) and for the insularity (*Island*) or the landlocked status of countries in the pair (*Landl*). Overall, the results for the time-invariant controls are economically meaningful in sign and size and highly statistically significant. With regard to the estimated coefficients for the economic integration agreements, again the dummies for both regional trade agreements and GATT/WTO have point estimates that are positive (0.578 and 0.365, respectively) and statistically significant at the 1 percent level of significance. However, in this specification the currency union dummy presents a counterintuitive sign. Anyway, unobserved bilateral heterogeneity and the likely endogeneity of economic integration agreements may be biasing the coefficient estimates (upwards or downwards).

Table 1 confirms the existence of a positive GATT/WTO effect on trade that has been previously found in most of the subsequent work to Rose's (2004) seminal contribution. However, as noted before, all these estimations may yield biased results since they do not account simultaneously for the previously discussed sources of bias in a single regression (hetoroskedastic residuals, zeros, endogeneity, unobserved bilateral heterogeneity and multilateral resistance). With the aim of dealing comprehensively with all these concerns, we estimate the gravity equation (1) with PPML. The results are displayed in column 1 of Table 2. The point estimates for regional trade agreements and currency unions are positive and statistically significant at the 1 percent level. However, an interesting result emerges with regard to our variable of interest since the

GATT/WTO effect vanishes once we include the full set of fixed effects in the PPML estimator.

In the remaining columns of Table 2, we re-examine the GATT/WTO effect on trade by different periods and groups of countries in order to test the robustness of our findings. In column 2, we investigate whether the trade effect over the GATT period (1948-1994) has been different from the trade effect over the WTO period (1995-2013). As we can see, the results do not reveal significant differences. In both cases, the estimated coefficients for GATT/WTO are not statistically significant at conventional levels. Later, we will further analyze the GATT/WTO effect for other alternative sub-periods using data for consecutive years.

Column 3 presents the results when we distinguish between early joiners (those countries that adhered to the GATT in the year of entry into force) and late joiners (those that joined the multilateral agreement in 1949 or later). To this end, we split the GATT/WTO dummy into two dummies: *GATTbothlatejoiners* (one for pairs of countries that joined the GATT after 1948); and *GATToneearlyonelate* (one for pairs including both kinds of countries).[19] Interestingly, the results show no GATT/WTO trade effects again.

In the last two columns of Table 2, we re-examine the effect of GATT/WTO across groups of countries with a standard classification criterion in this literature (industrialized versus developing countries).[20] In column 4, we disaggregate the GATT/WTO dummy into three dummies: one for industrialized country members (*GATTInd_Ind*), another for developing country members (*GATTDev_Dev*) and the

[19] It is worth noting that, since export data is available from 1948, the GATT trade effects between the 23 countries that joined the GATT in that year cannot be estimated because they are absorbed by the country-pair fixed effects.

[20] Several papers have addressed the GATT/WTO effect on trade distinguishing between industrial and developing countries with remarkably mixed results (Subramanian and Wei, 2007; Felbermayr and Kohler, 2010; Eicher and Henn, 2011; Dutt *et al*., 2013; Kohl, 2015; Bista, 2015; and Gil-Pareja *et al.* 2016). However, only do the last two papers take into account the group which each country in the pair belongs to (as we do here).

other for pairs combining industrial and developing country members (*GATTInd_Dev*). Column 5 further disaggregates the *GATTInd_Dev* dummy taking into account the direction of the export flows between members: from industrial countries to developing countries (*GATTIndExp_DevImp*) and from developing countries to industrial countries (*GATTDevExp_IndImp*). The results unequivocally reveal the absence of a positive GATT/WTO trade effect.

Table 3 presents the estimates for the same specifications of Table 2 when using data for consecutive years instead of data at five-year intervals. All the conclusions remain qualitatively unaltered. Both regional trade agreements and currency unions boost trade in the five specifications. However, the effect of GATT/WTO accession are estimated to be either nonexistent or even negative in three of the 12 point estimates reported.

In order to dig deeper into the impact of GATT/WTO on trade we further carry out the analysis by periods with other classification criteria. First, we restrict the sample period by rounds of trade negotiations (in a cumulative way). Second, we split the 66 years of sample period into six sub-periods with the same number of years. To this end, we use data for consecutive years (instead of data at five-year intervals) to guarantee the inclusion of the first and the last year of each period. The results when we confine the sample by rounds of trade negotiations are reported in Panel A of Table 4. The first period considered goes from 1948 to Dillon round (1961), the second one up to Kennedy round (1967), the third one up to Tokyo round (1979) and the fourth one up to Uruguay round (1994). It is remarkable that, in the four cases, the estimated coefficient of GATT/WTO dummy is never positive. Indeed, it is even negative and statistically significant (at least at the 10 percent level of significance) in three of the four cases.[21]

[21] This result is in line with Felbermayr and Kohler (2010), who show negative effects for the three time spans considered over the GATT period (1948-1994).

Panel B of Table 4 presents the results based on an alternative classification of the time periods. In particular, we split the 1948-2013 period into six periods of equal lengths of 11 years. [22] The results broadly confirm our previous findings. The estimated coefficients for the variable of interest are positive but statistically non-significant in three periods (1970-1980, 1992-2002 and 2003-2013), and negative and statistically significant at conventional significance levels in the other three periods (1948-1958, 1959-1969 and 1981-1991).

Furthermore, we investigate whether the change in the terms of accession for new entrants after the Uruguay Round (the obligation of a greater liberalization commitment for "new" developing countries that join the WTO since its creation than for the "old" developing countries that joined the GATT) has had an effect in the variable of interest. To this end, following Subramanian and Wei (2007), with cross-section data, and Gil-Pareja *et al*. (2016), with panel data for the period 1960-2008, we split the developing countries into two groups: those that were members before 1995 ("old members"); and those that become members since 1995 ("new members"). In particular, Table 5 displays the results when disaggregating the dummy variable GATT/WTO in four different ways.

Firstly, we disaggregate that dummy into three dummies depending on whether the importer is an industrialized country (*IndImp*), and old developing country (*OldDevImp*) or a new developing country (*NewDevImp*). We report the results in column 1. Next, we disaggregate the dummy variable GATT/WTO from the exporters' perspective disregarding the group, which the importer country belongs to (column 2). In column 3, we take into account the group which each trading partner (in the pair) belongs to (but not the direction of the trade flow). This involves splitting the

[22] This classification criterion follows Rose (2004) and Eicher and Henn (2011) that split their sample periods by decades. We have further split the sample period using different classification criteria and the results remain quantitatively and qualitatively unchanged. The results are available upon request.

GATT/WTO dummy into six dummies. As an example of the notation, *OldDev_NewDev* is a binary dummy variable that takes the value of one for pairs of member countries combining developing countries that joined the agreement before and after 1995. Finally, column 4 shows the results when we additionally take into account the direction of the bilateral export flows. For example, we define *OldDevExp_NewDevImp* as a variable that takes the value of one when the exporter is a developing country that joined the agreement before 1995 and the importer is a developing country that joined the agreement over the WTO period. The results confirm the absence of GATT/WTO effects on bilateral export flows. Only in one of the 21 cases reported in the table (exports from developing countries that joined WTO to industrial countries) the estimated coefficient of the variable of interest is positive and statistically significant at the 5 percent level.

So far, in all the specifications we have only considered the contemporaneous values of the variables for common membership in regional trade agreements, currency unions and the GATT/WTO. However, as Baier and Bergstrand (2007) noted, many agreements are "phased-in" over time (typically over 10 to 15 years), and terms-of-trade changes tend to have lagged effects on trade volumes. In order to account for these effects, we re-run the regression in column 1 of Table 2 including lags of the dummies *PTA*, *CU* and *GATT/WTO*. Columns 1 to 3 of Table 6 report the results when adding one, two and three lags for these variables, respectively. In order to see more easily the cumulative impact of the inclusion of lags, in these specifications we report the sum of the estimated coefficients from current and lagged values (denoted in Table 6 with the name of the variables without subscripts). We find that regional trade agreements have positive and statistically significant lagged effects increasing the point estimate from 0.183 without lags (column 1 of Table 2) to 0.244 (with one lag), 0.283 (with two lags) and 0.334 (with three lags). A similar pattern emerges for currency unions. In this case, the point estimates rise from 0.133, considering only the current effect, to 0.553 incorporating three lags of the variable in the regression. In both cases, the coefficient estimates have economically meaningful values. With regard to our variable of interest, the point estimates also increase with the inclusion of lags, but the estimated coefficient for the cumulative effect only reaches the statistical significance in the specification that includes three lags and simply to the 10 percent level of significance (column 3).

Moreover, columns 4 to 7 of Table 6 display the results when we further add one lead to the following four alternative specifications: without lags; with one lag; with two lags; and with three lags. This allows us testing for strict exogeneity of economic integration agreements (Wooldridge, 2010). Three comments are in order. Firstly, as before, the point estimates for regional trade agreements and currency unions are statistically significant and continuously raise from the specification without lags (column 4) to the specification with three lags (column 7). Secondly, this is also true for the *GATT/WTO* variable, but it does not reach the statistical significance at conventional levels in any case. Finally, despite accounting for endogeneity of economic integration agreements by including country-pair fixed effects in the regressions, the coefficient estimate for the lead of one of the three variables, regional trade agreements ($RTA_{ij,t+1}$), is positive and statistically significant at least at the 5 percent level of significance. The point estimate of this variable ranges from 0.066 to 0.079 in the four specifications suggesting some reverse causality. However, it is worth noting that the point estimates of the cumulative effects with lags for the *RTA* variable are somewhat larger in the specifications that include one lead (columns 5 to 7) than in those that do not include it (columns 1 to 3). Therefore, it seems that this econometric problem has a minor effect on the conclusions that arise when we do not include a lead of the *RTA* variable.[23]

Next, as an additional robustness check, we examine whether the evidence of no GATT/WTO effects still holds when we exclude from the regressions either the dummy variable for regional trade agreements, the dummy for currency unions or both at once. Before presenting the results, it should be stressed that a model that deletes one or more variables that are significant risks omission bias and inconsistency of the regression coefficients for the remaining economic integration agreements. However, this exercise is interesting here because all our previous results remain unaltered. For comparison purposes, column 1 of Table 7 reports again the results for the full specification, that is, the regression that includes the dummies for the three types of economic integration agreements *(RTA*, *CU*, and *GATT/WTO*). Column 2 presents the results when we exclude from the estimated regression the dummy for *RTA*. Regression in column 3 excludes the dummy for *CU*, whereas regression in column 4 excludes both. As we can see, the point estimate of the variable of interest hardly varies in a range that goes from -0.042 in the full specification to -0.070 in the specification that only includes the GATT/WTO dummy, and it is not statistically significant in any case. Moreover, the estimated coefficients for *RTA* and *CU* do not change either, even when we additionally exclude from the specification the GATT/WTO dummy variable (columns 5 and 6).

5. Conclusions

Rose's (2004) seminal paper prompted an intense debate on the effect of GATT/WTO on bilateral trade flows. This author strikingly documented the absence of

[23] For instance, in the regression with three lags (and no lead) the point estimate for the *RTA* variable is 0.334, which implies a cumulative effect of 39.7 percent [exp(0.334)-1]. By contrast, when we account for strict exogeneity by including the lead, the point estimate raises to 0.393 (and the corresponding cumulative effect to 48.1 percent). However, in this case there is evidence of a moderate "feedback effect" (8.0 percent) from trade to regional trade agreements.

GATT/WTO effects on trade, but much of the subsequent work has concluded that GATT/WTO has had trade enhancing effects. The empirical work addressing this question has progressively improved the econometric specifications in order to account for potential sources of bias. However, computational issues have conditioned the choice of estimator. The large datasets used in the estimation of GATT/WTO effects and/or difficulties to achieve convergence have precluded accounting simultaneously for unobserved bilateral heterogeneity (with country-pair fixed effects), for multilateral resistance terms (with exporter-time and importer-time fixed effects), as well as for heteroskedastic residuals and zero trade flows (with PPML).

This paper re-examines this issue taking advantage of recent econometric developments that allow us estimating structural gravity equations with PPML on a large dataset requiring to compute three types of high-dimensional fixed effects: exporter-time, importer-time and country-pair fixed effects. Our results are clearly supportive to Rose's (2004) findings. That is, in contrast to the trade-enhancing effect of both regional trade agreements and currency unions, GATT/WTO does not seem to have encouraged trade. In particular, we show that when we do not simultaneously account for all sources of estimation bias, the GATT/WTO effect on trade is positive. However, when we comprehensively account for all of these sources of bias our results contrast with conventional wisdom and the vast majority of previous empirical results: GATT/WTO accession does not generate statistically significant positive trade effects. Moreover, the results are robust across time periods and country groups using several alternative criteria of classification for both periods and groups of countries. These findings remain unchanged when we use data for consecutive years instead of data at five-year intervals. Finally, the results also hold when we take into account the GATT/WTO accession dynamics.

Noteworthy, our results do not deny the existence of some positive indirect effects of GATT/WTO on promoting trade, such as a generalized fall in trade barriers and more transparent, predictable and trade facilitating environment. These factors might have prompted regional trade agreements that seem to have boosted trade. Of course, these issues need further research.

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Table 1. OLS and Poisson results. Sample period 1948-2013 at five-year intervals.

Variables	(1) OLS with FE	(2) PML with CPFE	(3) PPML with CYFE
$LnGDP_{it}$		1.050	
		$(0.059)^{***}$	
$LnGDP_{jt}$		0.746	
		$(0.052)^{***}$	
Ln $Dist_{ij}$			-0.784
			$(0.014)^{***}$
$Cont_{ij}$			0.356
			$(0.032)^{***}$
$Lang_{ij}$			0.084
			$(0.028)^{***}$
$Colony_{ij}$			0.571
			$(0.038)^{***}$
$ComCount_{ij}$			1.682
			$(0.123)^{***}$
$Island_{ij}$			-0.083
			$(0.200)^{***}$
$Landl_{ij}$			-1.000
			$(0.113)^{***}$
RTA_{ijt}	0.399	0.309	0.578
	$(0.023)^{***}$	$(0.045)^{***}$	$(0.029)^{***}$
CU_{ijt}	0.417	0.344	-0.147
	$(0.049)^{***}$	$(0.040)^{***}$	$(0.048)^{***}$
$GATT/WTO_{ijt}$	0.162	0.224	0.365
	$(0.031)^{***}$	$(0.052)^{***}$	$(0.072)^{***}$
Time dummies	No	Yes	No
CYFE	Yes	No	Yes
CPFE	Yes	Yes	No
No observat.	152,406	130,671	155,951
R^2	-	-	0.90
Within R^2	0.87	-	-

Notes: The regressand in column 1 is the log of bilateral exports (lnX_{ijt}).. The regressand in columns 2 and 3 is the value of bilateral export flows (X_{ijt}). 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. Coefficient estimates for the fixed effects are not reported to save space.

Table 2. PPML estimation results. Sample period 1948-2013 at five-year intervals.

Variables	(1)	(2)	(3)	(4)	(5)
RTA_{ijt}	0.183	0.182	0.184	0.181	0.181
	(0.035)***	(0.035)***	(0.034)***	(0.034)***	(0.034)***
CU_{ijt}	0.133	0.132	0.129	0.130	0.130
	(0.043)***	(0.043)***	(0.043)***	(0.043)***	(0.043)***
$GATT/WTO_{ijt}$	-0.042				
	(0.077)				
$GATT\_1948\_1994_{ijt}$		-0.121			
		(0.090)			
$WTO\_1995\_2013_{ijt}$		0.114			
		(0.072)			
$GATTbothlatejoiners_{ijt}$			-0.123		
			(0.078)		
$GATToneearlyonelate_{ijt}$			-0.018		
			(0.079)		
$GATTInd\_Ind_{ijt}$				0.053	0.052
				(0.119)	(0.119)
$GATTDev\_Dev_{ijt}$				0.007	0.003
				(0.110)	(0.109)
$GATTInd\_Dev_{ijt}$				-0.048	
				(0.080)	
$GATTIndExp\_DevImp_{ijt}$					-0.075
					(0.072)
$GATTDevExp\_IndImp_{ijt}$					-0.030
					(0.106)
CYFE	Yes	Yes	Yes	Yes	Yes
CPFE	Yes	Yes	Yes	Yes	Yes
No observations	155,951	155,951	155,951	155,951	155,951

Notes: The regressand is the value of bilateral export flows. Robust standard errors (clustered by country-pairs) are in parentheses.* significant at 10%; ** significant at 5%; *** significant at 1%. The dummy variable *GATTbothlatejoiners* is one for pairs that joined GATT after 1948, and zero otherwise. *GATToneearlyonelate* is one for pairs combining one early joiner (GATT member in 1948) and one late joiner. CYFE indicates time-varying exporter and importer fixed effects. CPFE indicates country-pair fixed effects. Coefficient estimates for CYFE and CPFE are not reported for brevity.

Table 3. PPML estimation results. Sample period 1948-2013. Annual data for consecutive years.

Variables	(1)	(2)	(3)	(4)	(5)
RTA_{ijt}	0.197	0.196	0.198	0.196	0.196
	(0.038)***	(0.038)***	(0.036)***	(0.036)***	(0.036)***
CU_{ijt}	0.108	0.107	0.104	0.107	0.107
	(0.041)***	(0.041)***	(0.041)**	(0.041)***	(0.041)***
$GATT/WTO_{ijt}$	-0.110				
	(0.074)				
$GATT\_1948\_1994_{ijt}$		-0.163			
		(0.090)*			
$WTO\_1995\_2013_{ijt}$		0.044			
		(0.075)			
$GATTbothlatejoiners_{ijt}$			-0.207		
			(0.075)***		
$GATToneearlyonelate_{ijt}$			-0.082		
			(0.078)		
$GATTInd\_Ind_{ijt}$				-0.022	-0.023
				(0.097)	(0.096)
$GATTDev\_Dev_{ijt}$				-0.098	-0.100
				(0.103)	(0.103)
$GATTInd\_Dev_{ijt}$				-0.118	
				(0.076)	
$GATTIndExp\_DevImp_{ijt}$					-0.149
					(0.072)**
$GATTDevExp\_IndImp_{ijt}$					-0.095
					(0.113)
CYFE	Yes	Yes	Yes	Yes	Yes
CPFE	Yes	Yes	Yes	Yes	Yes
No observations	731,826	731,826	731,826	731,826	731,826

Notes: The regressand is the value of bilateral export flows. Robust standard errors (clustered by country-pairs) are in parentheses.* significant at 10%; ** significant at 5%; *** significant at 1%. The dummy variable *GATTbothlatejoiners* is one for pairs that joined GATT after 1948, and zero otherwise. *GATToneearlyonelate* is one for pairs combining one early joiner (GATT member in 1948) and one late joiner. CYFE indicates time-varying exporter and importer fixed effects. CPFE indicates country-pair fixed effects. Coefficient estimates for CYFE and CPFE are not reported for brevity.

Table 4. GATT/WTO effects by subperiods. Annual data for consecutive years.

	Panel A. By cumulative rounds of trade negotiations				Panel B. By 11-year periods					
Variable	(1) 1948-1961	(2) 1948-1967	(3) 1948-1979	(4) 1948-1994	(5) 1948-1958	(6) 1959-1969	(7) 1970-1980	(8) 1981-1991	(9) 1992-2002	(10) 2003-2013
RTA_{ijt}	0.129	0.327	0.323	0.426	0.028	0.365	0.260	0.221	0.025	0.019
	(0.036)***	(0.043)***	(0.043)***	(0.041)***	(0.046)	(0.047)***	(0.038)***	(0.035)***	(0.019)	(0.021)
CU_{ijt}	0.183	0.244	0.810	0.850	0.118	0.175	0.272	-0.106	-0.091	0.033
	(0.082)**	(0.067)***	(0.107)***	(0.092)***	(0.112)	(0.054)***	(0.038)***	(0.084)	(0.020)	(0.057)
$GATT/WTO_{ijt}$	-0.131	-0.137	-0.074	-0.226	-0.143	-0.111	0.094	-0.447	0.068	0.052
	(0.075)*	(0.059)**	(0.057)	(0.084)***	(0.073)**	(0.049)**	(0.068)	(0.079)***	(0.048)	(0.054)
CYFE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
CPFE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
No observ.	47,833	80,087	176,547	356,887	34,998	58,078	92,867	126,914	187,299	251,056

Notes: The regressand is the value of bilateral export flows. 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. Coefficient estimates for CYFE and CPFE are not reported for brevity.

Table 5. GATT versus WTO developing country members. Sample period 1948-2013 at five-year intervals.

Variables	(1)	(2)	(3)	(4)
RTA_{ijt}	0.183	0.182	0.188	0.188
	(0.034)***	(0.035)***	(0.032)***	(0.031)***
CU_{ijt}	0.134	0.132	0.132	0.132
	(0.043)***	(0.043)***	(0.043)***	(0.043)***
$IndImp_{ijt}$	-0.055			
	(0.098)			
$OldDevImp_{ijt}$	-0.092			
	(0.084)			
$NewDevImp_{ijt}$	0.042			
	(0.093)			
$IndExp_{ijt}$		-0.099		
		(0.081)		
$OldDevExp_{ijt}$		-0.068		
		(0.055)		
$NewDevEmp_{ijt}$		0.087		
		(0.067)		
Ind_Ind_{ijt}			-0.072	-0.073
			(0.126)	(0.124)
$OldDev\_OldDev_{ijt}$			0.162	0.193
			(0.162)	(0.155)
$NewDev\_NewDev_{ijt}$			0.157	0.156
			(0.096)	(0.096)
$OldDev\_NewDev_{ijt}$			0.006	
			(0.106)	
$OldDev\_Ind_{ijt}$			-0.264	
			(0.131)**	
$NewDev\_Ind_{ijt}$			0.171	
			(0.085)	
$OldDevExp\_NewDevImp_{ijt}$				0.056
				(0.140)
$NewDevExp\_OldDevImp_{ijt}$				-0.031
				(0.131)
$OldDevExp\_IndImp_{ijt}$				-0.365
				(0.151)**
$IndExp\_OldDevImp_{ijt}$				-0.119
				(0.136)
$NewDevExp\_IndImp_{ijt}$				0.220
				(0.097)**
$IndExp\_NewDevImp_{ijt}$				0.114
				(0.114)
CYFE	Yes	Yes	Yes	Yes
CPFE	Yes	Yes	Yes	Yes
No observat.	155,951	155,951	155,951	155,951

Notes: The regressand is the value of 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. Coefficient estimates for CYFE and CPFE are not reported for brevity.

Table 6. PPML estimation results with lags and leads. Sample period 1948-2013 at five-year intervals.

Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	1 Lag	2 Lags	3 Lags	1 Lead	1 Lag & 1 Lead	2 Lags & 1 Lead	3 Lags & 1 Lead
RTA	0.244	0.283	0.334	0.185	0.267	0.359	0.393
	(0.042)***	(0.046)***	(0.042)***	(0.034)***	(0.043)***	(0.042)***	(0.045)***
$RTA_{ij,t+1}$				0.066	0.070	0.079	0.077
				(0.029)**	(0.028)**	(0.027)***	(0.027)***
CU	0.147	0.215	0.553	0.141	0.174	0.656	0.668
	(0.047)***	(0.056)***	(0.082)***	(0.038)***	(0.046)***	(0.084)***	(0.093)***
$CU_{ij,t+1}$				-0.028	-0.030	-0.030	-0.046
				(0.042)	(0.040)	(0.037)	(0.036)
GATT/WTO	0.067	0.177	0.247	-0.074	0.064	0.155	0.218
	(0.106)	(0.120)	(0.133)*	(0.085)	(0.113)	(0.124)	(0.138)
$GATT/WTO_{ij,t+1}$				0.008	0.022	0.071	0.097
				(0.085)	(0.090)	(0.093)	(0.096)
CYFE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
CPFE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
No observations	153,152	149,961	146,078	133,830	131,031	127,840	123,957

Notes: The regressand is the value of bilateral export flows. Robust standard errors (clustered by country-pairs) are in parentheses.* significant at 10%; ** significant at 5%; *** significant at 1%. Regressions in columns 1 to 3 and 5 to 7 include, in addition to the current values of the dummies *RTA*, *CU* and *GATT/WTO*, one, two or three lags of these variables. In the specifications with lags, the table reports the sum of the estimated coefficients from current and lagged values for each variable using the "lincom" command in Stata. The cumulative effect of current and lagged variables is reported by the variable with no subscripts. CYFE indicates time-varying exporter and importer fixed effects. CPFE indicates country-pair fixed effects. Coefficient estimates for CYFE and CPFE are not reported for brevity.

Table 7. PPML estimation results. Sample period 1948-2013 at five-year intervals.

Variables	(1)	(2)	(3)	(4)	(5)	(6)
RTA	0.183		0.183		0.183	
	(0.035)***		(0.035)***		(0.035)***	
CU	0.133	0.132				0.133
	(0.043)***	(0.043)***				(0.043)***
GATT/WTO	-0.042	-0.060	-0.052	-0.070		
	(0.077)	(0.077)	(0.077)	(0.077)		
CYFE	Yes	Yes	Yes	Yes	Yes	Yes
CPFE	Yes	Yes	Yes	Yes	Yes	Yes
No observations	155,951	155,951	155,951	155,951	155,951	155,951

Notes: The regressand is the value of bilateral export flows. 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. Coefficient estimates for CYFE and CPFE are not reported for brevity.