

WTO Trade Effects and Identification Problems: Why Knowing The Structural Properties of WTO Memberships Matters?

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Abstract

Since Rose's (2004) striking finding of negligible WTO trade effects, numerous studies have attempted to solve the so-called WTO puzzle. These studies adopt novel model specifications to control for potential sources of bias, but often lead to conflicting results. Multilateral resistance terms (MRTs), unobserved country-pair heterogeneity (UCPH) and heteroskedastic errors in log-linear model are considered the most crucial controls. What has gone unnoticed, however, is that the first two controls lead to identification problems in the estimation of WTO trade effects. We show that controlling for MRTs leads to near-perfect multicollinearity because of a structural relationship between the variables that measure the GATT/WTO membership statuses of any country-pairs. Also because of this structural relationship, accounting for UCPH using country-pair fixed effects (CPFEs) could reduce the number of observations that contribute to the identification of WTO trade effects by more than 98%. These identification problems make the estimates of WTO trade effects very imprecise and sensitive to model specifications, partly explaining the diverse results in the literature. We propose a two-stage method that avoids these identification problems

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and finds a strong positive WTO trade effect. It is found that bilateral trade flows between two countries increase even when only one of them joins the WTO, but the effect is much bigger when both are WTO members. The positive findings from the two-stage method are robust to various model specifications as long as both MRTs and CPFEs are controlled for.

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1 Introduction

Since Rose (2004) first documented the surprising results that GATT/WTO membership (hereafter we use “WTO” to represent “GATT/WTO” for brevity) does not have statistically discernible impacts on bilateral trade flows, much literature discussing the so-called WTO puzzle has emerged. A number of studies have reexamined Rose’s findings. Amongst them, Tomz, Goldstein, and Rivers (2007), Subramanian and Wei (2007), Liu (2009), Chang and Lee (2011) and Konya, Matyas, and Harris (2011) reject the puzzle, but Eicher and Henn (2011) and Roy (2011) reconfirm it. The fact that evidence supporting the effect of such a large policy coordination enterprise is so fragile is sufficient enough to cast doubt on those economic theories giving unequivocal conclusion regarding the outcome of trade liberalization, and to cause frustration amongst those advocating the WTO membership. As such, there is a lot at stake in solving the WTO puzzle.

The recent literature on the empirical gravity equation of bilateral trade flows emphasizes the need to account for potential biases due to multilateral resistance terms (MRTs), unobserved country-pair heterogeneity (UCPH) and heteroskedasticity. Accounting for MRTs, which are implicit price indexes, is important because of their non-linear effects on bilateral trade flows between countries (Anderson and van Wincoop, 2003).¹ A reduction of trade costs due to the WTO membership will affect a country’s MRTs; as such, not properly controlling for MRTs in the estimation of WTO trade effects will cause omitted variable bias (OVB). The deployment of country-pair fixed effects (CPFEs) to account for UCPH is also emphasized in the literature because the WTO membership status of a given country-pair may be associated with their unobserved bilateral affinities.² In addition, Santos Silva and Tenreyro (2006) show a large bias in a gravity model with log-linear specification due to heteroskedastic errors and recommends the Poisson Pseudo Maximum Likelihood (PPML) method to address it. In fact, numerous studies find that additional controls for MRTs, UCPH, or heteroskedasticity can dramatically change the estimates of WTO trade effects and lead to conflicting conclusions.

¹Also see Feenstra 2004; Subramanian and Wei 2007; Eicher and Henn 2011; Roy 2011; Eicher, Henn, and Papageorgiou 2012.

²See Cheng and Wall (2005); Baldwin and Taglioni (2006); Baier and Bergstrand (2007); Eicher and Henn (2011).

For instance, Liu (2009) reverses Rose’s findings by accounting for UCPH, heteroskedasticity and including zero observations, but the additional control for MRTs in Roy (2011) once again resurrects the puzzle. Dutt, Mihov, and Van Zandt (2011) also reveal that the sign of WTO trade effects flips as their estimation models progressively control for UCPH and MRTs. Eicher and Henn (2011) overturn Subramanian and Wei’s (2007) findings once additionally accounting for UCPH and individual PTAs. Table 1 summarizes the diverse results in previous studies on WTO trade effects.

All but one of the studies in Table 1 use the gravity model. The exceptional one, Chang and Lee (2011), obtains positive WTO trade effects using Rose’s data by applying a more flexible matching method. The matching method is a non-parametric technique that can avoid biases from parametric specifications. Most of the recent studies in Table 1 consider UCPH as unobserved factors which, however, can not be easily incorporated into the matching method. Moreover, since parametric methods are more popular in this literature, using the same approach allows direct comparison of our findings with a larger number of studies. Therefore, in this paper we restrict our analyses to parametric models.

We find that two hitherto unnoticed identification problems arise when controlling for MRTs and UCPH in parametric models, partly explaining the diverse results in previous studies. These problems are caused by the structural relationship between the two variables used throughout the literature to indicate whether a country-pair has single or joint WTO membership (with no membership being the baseline) and result in very imprecise and specification-sensitive estimates. These problems are evident as they will not occur if one considers the effect of joint WTO membership only.³ In doing so, those studies use country-pairs with either no membership or a single membership as the control group, and those with joint memberships as the treatment group. However, this specification could yield misleading conclusions on the WTO trade effects if there are significant differences in bilateral trade flows across no membership country-pairs and their single membership counterparts. For instance, even though the WTO system is based on the reciprocity principle, some WTO members

³Examples include Subramanian and Wei (2007); Felbermayr and Kohler (2010); Eicher and Henn (2011); Dutt, Mihov, and Van Zandt (2011).

appear to extend the most-favored-nation (MFN) principle to non-member countries (Rose, 2004); as such, WTO trade effects for single membership could be not different from joint memberships.⁴

Studies that consider both single and joint WTO memberships (e.g. Rose, 2004, 2005; Tomz, Goldstein, and Rivers, 2007; Liu, 2009; Roy, 2011), are confronted with two very different identification problems when accounting for MRTs and UCPH. In this paper we demonstrate that, although accounting for MRTs can mitigate OVB, it creates near-perfect multicollinearity. As a result, despite the fact that unbiased point estimates of WTO trade effects may be very large in economic terms, it can still be statistically insignificant because of the equally large standard error estimates associated with them. The near-perfect multicollinearity also results in very different estimates with even small changes in data or specification. For instance, when Roy (2011) uses the same dataset but additionally accounts for MRTs, the originally large significant and positive coefficient estimates turn into large negative coefficients that are not significant. Secondly, we find that although controlling for UCPH is important to reduce OVB, it could reduce the number of observations contributing to the identification of WTO trade effects by as much as 98.4%. We will show that the problems of multicollinearity and the small number of effective observations are both attributed to certain structural properties of the two WTO membership variables.

In this paper we propose a two-stage method for estimating WTO trade effects that overcomes the aforementioned identification problems. The first stage for WTO accession for any country-pairs, starting from no membership, is for only one of the two countries to accede to the WTO. The second stage is then when the other country accedes as well. In our dataset, only 1.6% of country-pairs change directly from no membership to joint memberships. This approach allows us to estimate the WTO trade effects in each of the two stages and compare single and joint memberships effects. More importantly, by separating the two stages, it avoids the multicollinearity and small number of observations problems even when controlling for MRTs and UCPH.

⁴There can be other reasons why there are differences between no membership and single membership: On the one hand, a single WTO membership may have negative effects on bilateral trade if a member country's trade is diverted from a non-member to a member. On the other hand, a single WTO membership may have positive effects if, first, a member applies the WTO rules (e.g. restrictions on subsidies) to non-member countries, and second, firms in a member country become more productive and export more to non-member countries.

The new method provides evidence of strong positive WTO trade effects at both stages. Both single and joint memberships are found to have statistically and economically significant effects on bilateral trade flows, but the effects for joint membership are much greater. The results hold for both developing and developed countries as well as for both formal and informal WTO membership. However, the effects vary over different time periods. For instance, until the 1990s, only single membership is found to have positive trade effects. This finding implies that only testing for joint memberships, as done in some previous studies, may result in misleading conclusions. Chang and Lee (2011) are concerned with potential bias in parametric linear gravity models due to model misspecification. To avoid this pitfall, they adopt a more flexible matching method and obtain positive WTO trade effects. Our results indicate that this discrepancy, as well as the sensitive estimates in the parametric gravity models, may be as a result of other factors such as identification problems rather than the model misspecification.

This paper is also one of the first studies in the WTO trade effects literature that take into account all the biases arising from MRTs, UCPH, a large number of zero observations and heteroskedasticity (see Liu (2009) and Eicher and Henn (2011) for the summary).⁵ We find that coefficient estimates vary widely according to model specifications – including zero observations or not, using log-linear or the Poisson models, and accounting for UCPH or not – even in the absence of multicollinearity, albeit circumventing multicollinearity improves precision significantly. This implies that UCPH, heteroskedasticity, zero observations and MRTs all need to be properly accounted for in the estimation.

The rest of the paper is organized as follows. Section 2 explains the econometric methodology and data. Section 3 reports the main results and section 4 shows various extensions and sensitivity analyses. The last section concludes with a discussion of the implications.

⁵Liu (2009) considers all these elements; however the remoteness variable that used to control for MRTs is rather ad hoc as it is not based on the theory of MRTs as originally developed in Anderson and van Wincoop (2003).

2 Methodology and Data

2.1 Gravity model

Our empirical tool is the gravity model. A log-linear gravity model based on the general equilibrium framework suggested by Anderson and van Wincoop (2003) can be expressed as:

$$\begin{aligned} \ln(T_{ijt}) = & \alpha_0 + \alpha_1 \ln Y_{it} + \alpha_2 \ln Y_{jt} + \alpha_3 \ln y_{it} + \alpha_4 \ln y_{jt} + \mathbf{X}_{ijt} \cdot \boldsymbol{\beta} \\ & + \gamma_1 \text{Onein}_{ijt} + \gamma_2 \text{Bothin}_{ijt} + \mu_t + \omega_{ij} - \ln P_{it}^{1-\sigma} - \ln P_{jt}^{1-\sigma} + \epsilon_{ijt} \end{aligned} \quad (1)$$

subject to the N nonlinear market-equilibrium conditions

$$\ln P_{it}^{1-\sigma} = \sum_{k=1}^N \ln P_{kt}^{1-\sigma} (Y_{kt}/Y_{Wt}) e^{\mathbf{X}_{ijt} \cdot \boldsymbol{\beta} + \gamma_1 \text{Onein}_{ijt} + \gamma_2 \text{Bothin}_{ijt}}, i = 1, \dots, N \quad (2)$$

where *Onein* is a dummy variable that takes a value of one if either importer i or exporter j was a WTO member in year t and zero otherwise; *Bothin* is another dummy variable that takes a value of one if both i and j were WTO members in year t and zero otherwise; Y is real GDP measured in purchasing power parity (PPP) terms; y is real per capita GDP (PPP); $\mathbf{X}$ is a vector of trade cost variables to be detailed in the data section; μ_t represents any unobserved global trend in trade and aggregate shocks in each year; ω_{ij} is country-pair heterogeneity; $P_k^{1-\sigma}$ is a MRT; P_k is the price level of country k ; σ is the elasticity of substitution in consumption; and Y_{Wt} is world GDP at time t .

As MRTs are unobservable, Anderson and van Wincoop (2003) and Feenstra (2004) suggest to use country-time fixed effects (CTFEs) as a proxy for them in order to obtain unbiased estimates. However, when the dataset is large, it is computationally challenging to estimate the gravity model with CTFEs using the least square dummy variable (LSDV) method due to the large number of dummy variables involved. For instance, our dataset has 209 countries and 54 periods (1948-2001), applying the LSDV method to it could entail 21,736 dummy variables in total. To avoid this difficulty, previous studies

in this literature either limit the period of the dataset (e.g. Magee, 2008 and Roy, 2011), significantly reduce observations (e.g. Eicher and Henn, 2011), or do not control for CTFEs (e.g. Liu, 2009).

An alternative to the CTFEs approach is to use a first-order log-linear Taylor expansion of the MRTs as suggested by Baier and Bergstrand (2009a) (hereafter BB). They shows that (1) and (2) can be approximated as

$$\begin{aligned} \ln(T_{ijt}) = & \alpha_0 + \alpha_1 \ln Y_{it} + \alpha_2 \ln Y_{jt} + \alpha_3 \ln y_{it} + \alpha_4 \ln y_{jt} + (\mathbf{X}_{ijt} - MR\mathbf{X}_{ijt}) \cdot \beta \\ & + \gamma_1 (Onein_{ijt} - MROnein_{ijt}) + \gamma_2 (Bothin_{ijt} - MRBothin_{ijt}) + \mu_t + \omega_{ij} + \epsilon_{ijt}, \quad (3) \end{aligned}$$

and

$$MRz_{ijt} = \sum_{k=1}^N \left(\frac{Y_{kt}}{Y_{Wt}} \right) z_{ikt} + \sum_{m=1}^N \left(\frac{Y_{mt}}{Y_{Wt}} \right) z_{jmt} - \sum_{k=1}^N \sum_{m=1}^N \left(\frac{Y_{kt}}{Y_{Wt}} \right) \left(\frac{Y_{mt}}{Y_{Wt}} \right) z_{kmt} \quad (4)$$

where MRz is a linearized MRT, with $z_{ijt} \in \mathbf{Z}_{ijt} \equiv (\mathbf{X}_{ijt}, Onein_{ijt}, Bothin_{ijt})$. In other words, the method requires each of the ijt -varying variables to be “deMRTted” (as akin to “demeaned”) when entering the regression.⁶ WTO trade effects are measured by $\hat{\gamma}_1$ and $\hat{\gamma}_2$. The linearized MRTs are composed of three terms: multilateral resistance between the importing country and all countries in the world; multilateral resistance between the exporting country and all countries in the world, and; multilateral resistance amongst all countries in the world. The first two terms increase multilateral resistance between a country pair because, other things being equal, more of single or joint memberships between country i or j and the rest of world make trade between i and j more difficult, whereas the last term decreases multilateral resistance because more of single or joint memberships amongst

⁶Baier and Bergstrand (2009b) apply their method to estimate the border effect on trade between the US and Canada using a cross-sectional dataset. Our paper extends it to a large scale panel dataset and consider all the independent variables except GDP and income as sources of trade friction. that contribute to multilateral resistance. Silva and Nelson (2012) and Portugal-Perez and Wilson (2012) also apply the BB method for a large panel dataset in the gravity equation. Chang and Lee (2011) use this method to obtain MRTs as one of covariates in their matching framework.

countries in the world make trade between i and j easier.

Santos Silva and Tenreyro (2006) provide Monte Carlo simulation evidence that, even if a gravity model is correctly specified, heteroskedastic errors can cause severe bias on the OLS coefficient estimates in the log-linear model. They advocate using the Poisson Pseudo Maximum Likelihood (PPML) estimator which is consistent as long as the conditional mean of (5) is correctly specified.

$$E(T_{ijt}|\mathbf{W}_{ijt}, \mu_t, \omega_{ij}, P_{it}, P_{jt}) = \exp(\alpha_0 + \alpha_1 \ln Y_{it} + \alpha_2 \ln Y_{jt} + \alpha_3 \ln y_{it} + \alpha_4 \ln y_{jt} + \mathbf{X}_{ijt} \cdot \beta) \quad (5)$$

$$+ \gamma_1 \text{Onein}_{ijt} + \gamma_2 \text{Bothin}_{ijt} + \mu_t + \omega_{ij} - \ln P_{it}^{1-\sigma} - \ln P_{jt}^{1-\sigma}$$

where $\mathbf{W}_{ijt}$ is a vector of all observed variables.

Subsequently, the Conditional PPML (CPPML) method, which accounts for UCPH by conditioning on the sum of trade flows over time, is widely used in the literature (e.g. Magee, 2008; Liu, 2009; Konya, Matyas, and Harris, 2011). The CPPML, however, often encounters convergence problems when there are a large number of explanatory variables, making the CTFEs approach to MRTs not feasible. Considering this, the CPPML method with BB's linearized MRTs is our preferred approach. As long as following conditional expectation for trade flows is correctly specified, the CPPML estimates for γ 's are consistent:⁷

$$E(T_{ijt}|\mathbf{W}_{ijt}, \sum_{t=1} T_{ijt}) = \exp(\alpha_0 + \alpha_1 \ln Y_{it} + \alpha_2 \ln Y_{jt} + \alpha_3 \ln y_{it} + \alpha_4 \ln y_{jt} + (\mathbf{X}_{ijt} - MR\mathbf{X}_{ijt}) \cdot \beta)$$

$$+ \gamma_1 (\text{Onein}_{ijt} - MR\text{Onein}_{ijt}) + \gamma_2 (\text{Bothin}_{ijt} - MR\text{Bothin}_{ijt}) \quad (6)$$

We also report the estimates from two log-linear models with and without zero observations for comparison.⁸

⁷The BB's terms do not directly account for import and export country-time specific shocks unlike CTFEs. However, Adam and Cobham (2007) show that the estimates obtained from two methods are virtually identical.

⁸Helpman, Melitz, and Rubinstein (2008) incorporate firm heterogeneity to account for zero observations in the gravity equation. Their procedure takes into account sample selection assuming the fixed costs of exporting firms. However, we

2.2 Data

The main dataset is obtained from Liu (2009). It covers 209 countries from 1948 to 2001. The GDP and population data are obtained from the PWT6.1, PWT5.6, WDI2003, Maddison Historical Statistics, the IMF International Financial Statistics (IFS) and the United Nations Statistical Yearbooks (UNSYB). The US consumer price index is used to convert these GDP measures into 1995 real dollar terms. All GDP data used in this paper, except those from the IFS, are measured by purchasing power parity (PPP) methods. The PWT6.1 dataset is taken as the base source for GDP data and any missing data are filled in using data from other datasets after being multiplied by a ratio calculated from the overlapped GDP data in the two datasets. WTO formal membership and regional trade agreement (RTA-free trade agreement and customs union) data are obtained from the WTO website.

The vector of trade cost variables $\mathbf{X}$ includes distance (the great circle distance between i and j), area (the geographic area of a country), landlockedness (the number of landlocked nations in a pair (0, 1, or 2)), island (the number of island nations in a pair (0, 1, or 2)), common language (dummy equals to one if i and j share a common language), common religion (dummy equals to one if i and j share a common religion), colony (dummy equals to one if i has ever been a colony of j), colonizer (dummy equals to one if i has ever been a colonizer of j), current colony (dummy equals to one if i is a colony of j in year t), current colonizer (dummy equals to one if i is a colonizer of j in year t), common colony (dummy equals to one if i and j have ever been colonized by the same colonizer), hostility (the military conflict intensity between i and j), alliance (dummy equals to one if i and j were in a formal alliance in year t), currency unions (dummy equals to one if i and j used the same currency in year t), regional trade agreement (dummy equals to one if i and j belonged to the same free trade agreement or customs union in year t), and general system of preferences (GSP-dummy equals to one if i (j) offered GSP to j (i) in year t). Currency Union data are from Glick and Rose (2001). Generalized

have encountered two problems in the application of this method. Firstly, we cannot control for CPFEs in the first-stage Probit estimation as the method to do so has not yet developed in the econometric literature and it is computationally challenging to run a regression using the LSDV method. Secondly, finding a proper exclusion restriction that contains enough time variations for over 50 years is extremely difficult, and misspecification and poor choice of exclusion restriction are likely to increase bias.

System of Preferences (GSP) data are mainly from the UN publications, and some changes are made to the original data according to specific government publications on GSP. Latitudes, longitudes, areas, land contiguity, languages, religions, landlocked and island status are obtained from the 2003 CIA Fact Book. The great circle distances are constructed from the latitudes and longitudes of country-pairs.

Besides *de jure* WTO membership, we also consider *de facto* WTO membership, which includes both member and non-member participants of WTO as suggested by Tomz, Goldstein, and Rivers (2007) as robustness check.⁹

3 Main Results

3.1 Preliminary results

Table 2 reports the estimates for the *Onein* and *Bothin* variables using different model specifications. In the first two columns the log-linear model is used, and in the last two columns the (conditional) Poisson model is used. For each of the log-linear and the (conditional) Poisson models, we have tested for different sets of controls, namely: time fixed effects (TFEs), CPFEs, and MRTs. This gives us in total 14 sets of results, as two of the (C)PPML estimations do not converge.

The estimates for *Onein* and *Bothin* vary widely across different specifications, some are highly significant but of an unexpected negative sign (e.g. -0.20 for *Onein* in the first row). Even for the same log-linear model, there are huge differences between estimations with and without zero observations. The log-linear model with zero observations consistently yields significant, positive estimates for both variables; however, many of the estimates are implausibly large such as 10.74 for *Onein* and 20.84 for *Bothin* in the second column. These hugely diverse findings echo the fragility of WTO trade effects estimates in the literature.

The results of controlling for TFEs, CPFEs and MRTs are worth discussing more because, as this set of controls is most comprehensive, other things equal, the bias of the associated estimates should

⁹The data used are available at http://www.stanford.edu/~tomz/pubs/TGR_AER2007.zip.

be the smallest. For this set of controls, we also report the estimates of *RTA* for comparison, because a regional trade agreement reduces the trade barriers between two countries similar to WTO membership (especially for *Bothin*). It can be seen that there are no qualitative differences between the estimates of *RTA* across the log-linear and the PPML estimations in terms of sign and significance, and that the quantitative differences are also much smaller in comparison. As mentioned in the introduction, the current gravity equation literature emphasizes on controlling for UCPH, MRTs, zero observations and heteroskedasticity in the log-linear model. Because the PPML method with the full set of controls and zero observations consider them all, it is the preferred one amongst all estimations in Table 1. According to the CPPML estimates with zero observations, *RTA* will increase bilateral trade by about 48% ($= (e^{0.39} - 1) * 100$), which is similar to many findings in the literature (e.g. see Magee, 2008; Eicher, Henn, and Papageorgiou, 2012). On the other hand, WTO membership has no statistically discernible effects on bilateral trade flows, which runs against the expectation and thus reconfirms the WTO puzzle.

When we look more closely, the coefficients for *Onein* and *Bothin* in the preferred CPPML estimation in Table 1 are in fact of the correct signs and plausible magnitudes. The point estimate of 1.03 for *Bothin* implies that the joint WTO memberships for two countries are expected to increase their bilateral trade by 180% compared to the baseline that neither of them is a WTO member. Judging by this figure alone, WTO trade effects are incredibly large. However, at the same time the associated standard error is estimated to be 2.01, meaning that the point estimate for *Bothin* had to be greater than 5,000% to be statistically significant at the 5% significance level. The estimate for *Onein* is equally imprecise. In other words, despite the number of observations being over one million, the estimation is still not precise enough for the two variables. This seems to indicate that the WTO puzzle may be attributed to factors beyond the neglect of UCPH, MRTs, zero observations, and heteroskedasticity in the log-linear model.

3.2 Identification problems: multicollinearity

One of the most noticeable results from Table 1 is that the estimated standard errors for WTO trade effects increase when MRTs are accounted for. For instance, the estimated standard error for the joint WTO membership effect increases more than tenfold with the control of MRTs in the CPPML method. Multicollinearity between the two WTO membership variables is a primary suspect for this escalation.

Table 3 shows the correlation coefficient between the two WTO membership variables from the 1950s to the 1990s. Unconditional correlation reported in the first row is always negative and significant. It is attributed to a structural property of the WTO membership variables. By construction, for any dyad observation *Onein* and *Bothin* can have only possible combinations of values: $\{Onein = 0, Bothin = 0\}$ for no membership, $\{Onein = 1, Bothin = 0\}$ for a single membership, or $\{Onein = 0, Bothin = 1\}$ for joint memberships. Because around 75% of the observations in our sample are of either a single membership or joint memberships the correlation between the two membership variables is high and negative. In addition, as the number of countries acceding to WTO only increases over time,¹⁰ the share of country-pairs with no membership will get smaller and smaller, pushing the correlation between the two membership variables towards -1 .

In actual estimations it is the conditional correlation that matters. The second to fifth rows of the Table 3 reports the correlation coefficients conditional on various combinations of TFEs, CPFES, and MRTs (or CTFEs).¹¹ Although the correlation coefficients conditional on TFEs and CPFES are higher than the unconditional one (in absolute terms) in some cases, the figures do not seem to be large enough to cause severe multicollinearity. However, once MRTs (or CTFEs) are controlled for, the correlation suddenly become close to unity for every decade including the very first one. That is, controlling for MRTs (or CTFEs) can magnify the already high correlation between the two WTO membership variables to near-perfect correlation.

¹⁰In the history of WTO, there was only one incidence of a member – China – withdrawing its membership. However, China applied and readmitted into the WTO in subsequent years.

¹¹The correlations conditioning on TFEs and CTFEs are obtained using the regression residuals of the WTO membership variables conditioning on dummy variables. To obtain the correlations conditioning on CPFES and MRTs, we use the demeaned and “deMRTted” WTO membership variables, respectively. Because of the data size constraint of the STATA program, we are not able to obtain the correlation coefficient for the whole sample when CTFEs are controlled for.

Figure 1 illustrates how the control of MRTs using the BB method magnifies the correlation between the two membership variables. Panel (a) is a scatter plot of *Bothin* against *Onein*. They have only three possible combinations of values, $\{Onein = 0, Bothin = 0\}$, $\{Onein = 1, Bothin = 0\}$ and $\{Onein = 0, Bothin = 1\}$. Because many observations have identical values, the plot is depicted as a density distribution plot.¹² Each short line segment (petal) of an orange circle represents 44,617 observations. The fitted value line shows that the correlation of the two variables is negative but far from -1. Panel (b) is a scatter plot of *MRBothin* against *MROnein*. A small blue circle represents one observation, each petal of a light green circle also represents one observation, and one petal of an orange circle represents 13,392 observations. The density is very high around the negative-slope fitted value line. Lastly, panel (c) provides a scatter plot of $(Bothin - MRBothin)$ against $(Onein - MROnein)$. Again, the majority of observations are located along the fitted value line, resulting in near-perfect correlation between the two variables. Overall, Figure 1 illustrates that, because there exists a high and negative correlation between *MROnein* and *MRBothin*, the “deMRTing” process in the BB method leads to a dramatic increase in the correlation between the transformed WTO membership variables. A more detailed explanation for why (instead of how) the control of MRTs increases correlation between the two membership variables is provided in the appendix.

3.3 Identification problems: a small number of observations

Table 4 shows the distribution of WTO membership amongst all country-pairs in our dataset according to whether they have changed their WTO membership status. Amongst 35,532 pair of countries, 29.3% (6.3% + 18.7% + 4.3%) have never changed their WTO membership status throughout the sample period. For the remaining 70.7%, roughly one third (22.1%) have changed from no membership to single membership at some point and remained like that till the end of the sample period, and two third (47%) have changed twice from no membership to single membership and then further to joint memberships by the end of the period. A very small proportion (1.6%) changed from no membership

¹²See Dupont and Plummer (2005) for more detailed explanations on the density distribution plot.

directly to joint memberships. These figures reveal that, for country-pairs that do experience changes in WTO membership, it is almost always *one membership at a time*. If we disaggregate the data by decades, for example, in the 1970s, only 0.1% (4 out of 34, 571 country-pairs) changed from no membership to joint memberships. These data characteristics on membership status has three important implications. Firstly, our proposed two-stage method should be viewed as more appropriate considering the data structure. Second, models that combine those country-pairs of no membership with those of single membership as the reference groups, throw away 45.6% ($22.1\% + 47\%/2$) of the variation in membership status. Third, these data structure could affect the identification of joint WTO trade effects when CPFEs are controlled for. For example, in the traditional approach as in equation (1), the magnitude of γ_2 is measured by the effect of a change in *Bothin* on $\ln(T)$, conditional on all other variables being kept constant, including *Onein*. The identification of γ_2 can be based on (i) a cross sectional comparison of joint membership country-pairs with no membership country-pairs, and (ii) a temporal comparison for country-pairs that change from either no membership or single membership to joint memberships. However, the first part of this identification using between-variations of country-pairs, is eliminated when CPFEs are controlled for. In addition, if *Onein* is also held fixed, then the identification relies only on within-variations of each of the 1.6% country-pairs that change from no membership directly to joint memberships. This means that merely 552 ($35,532 \times 0.016$) country pairs have contributed to the identification of γ_2 . Given the findings are based on such a small fraction of the sample, it may not be appropriate to draw inferences from them for the wider population.

3.4 New results using the two-stage method

To address the multicollinearity problem and retain as many variations in the data for identification as possible, we propose a two-stage method in the estimation of WTO trade effects. The first stage estimates the effects from no membership to single membership, and the second stage estimates the effects from single membership to joint memberships. Let us define another dummy variable

$Nonein_{ijt}$ that takes a value of one if both i and j are not WTO member at time t and zero otherwise. The identification of the first-stage effect is based on comparing $Onein = 1$ observations with $Nonein = 1$ observations as the reference group, and $Bothin = 1$ observations are excluded. The second-stage effect is based on comparing $Bothin = 1$ observations with $Onein = 1$ observations as the reference group, and $Nonein = 1$ observations are excluded. This is in contrast to the estimation strategy in Table 2 where both single and joint WTO memberships effects are estimated simultaneously based on comparing $Onein = 1$ or $Bothin = 1$ observations with $Nonein = 1$ observations as the reference group. Under this two-stage estimation strategy, the multicollinearity problem is solved because the two WTO membership variables do not simultaneously enter the regression models. Furthermore, the change of the reference group can enhance the precision of γ_2 estimation because the within variations of 47% country-pairs are now used for its identification as compared to the 1.6% in the traditional approach.

The results of the stage-wise regressions are reported in Table 5. Although the results are represented in a similar layout as Table 2, the coefficients for $Onein$ and $Bothin$ are now estimated separately in different regressions. We focus on the bottom panel where TFEs, CPFes, and MRTs are controlled for. The results for RTA are also presented to contrast the stability of the RTA effect estimates to the fragility of WTO effects estimates. All the coefficients in the panel are significant at the 1%, 5% or 10% level, and have expected positive signs except in one case (-0.01) where the coefficient of $Onein$ is not significantly different from zero. The coefficient estimates for RTA are comparable to those in Table 2. Regarding the CPPML estimation with zero observations, the coefficients for $Onein$ and $Bothin$ are significant at the 5% and 1% level respectively, in great contrast to the highly insignificant results in Table 2. According to these estimates, the acquisition of the first WTO membership for a country-pair, other things being equal, is expected to increase their bilateral trade by 15%, while the acquisition of the second WTO membership increase their bilateral trade by another 153%. In magnitude, the second-stage effect is 10 times greater than the first-stage effect.

For the rest of the table, a number of findings are worth highlighting. Firstly, the standard errors for the coefficient of $Bothin$ are reduced in all specifications as compared to those in Table 2. In the

cases where MRTs are controlled for, the reduction of standard errors are particularly dramatic, from the range of 0.46-2.01 in Table 2 to 0.06-0.25 in Table 5. In fact, in those cases, the standard errors for the coefficient of *Onein* are also substantially reduced from the range of 0.23-0.99 to 0.02-0.07. The precision of the estimates for the single and joint WTO membership effects significantly improves in the two-stage method. Secondly, in the cases where MRTs are controlled for, the coefficient estimates in Table 5 are of much more plausible magnitudes than those in Table 2. Thirdly, the coefficient estimates for *Onein* and *Bothin* vary widely, as in Table 2, even with the two-stage method. It indicates that although the two-stage method directly improves the estimation precision significantly, it cannot remove the biases from UCPH, zero observations and heteroskedasticity. This means the two-stage method and proper controls for UCPH, MRTs, heteroskedasticity and zero observations should be considered all together.

4 Extensions

In this section we conduct several extensions and examine if WTO trade effects are sensitive to (i) the expansion of the membership definition; (ii) whether new members are importing or exporting countries; (iii) whether new members are developing or industrial countries; and (iv) using different decade subsamples. We continue to use the two-stage approach and account for UCPH, MRTs and zero observations.

4.1 *De facto* WTO trade effects

We consider a more expanded definition of WTO membership as suggested by Tomz, Goldstein, and Rivers (2007) that additionally includes three categories of countries that follow the WTO practice without acquiring the formal membership. These nonmember participating countries typically are colonies of WTO member countries, *de facto* members, and provisional members. The estimation results with the two-stage method are reported in Table 6. The qualitative results of Table 6 are the

same as those in Table 5, supporting the previous finding that positive and economically significant WTO effects are not sensitive to the definition of WTO membership.

4.2 Importer and exporter WTO trade effects

While the WTO has prohibited the use of industrial policies such as export subsidies, one of the most important agendas of the organization has been the systematic lowering of import tariffs. Given this, one may think that the positive WTO trade effects for single membership country-pairs in Table 5 would be primarily attributed to the positive effects from the importing country's membership dominating the potential negative effects from the exporting country's membership. Accordingly, we can examine distinct WTO trade effects depending on whether new members are importing or exporting countries, especially in the first stage. The results are shown in Table 7. The way Table 7 reports the estimates is slightly different from that of Tables 5 and 6 because we now distinguish between importing and exporting countries. If a new member is an importer (exporter), the first stage regression estimates the effect of changing from being in a no membership country pairing ($Nonein = 1$) to the importing (exporting) country having a membership ($Impoterin(Exporterin) = 1$), and in the second stage the regression estimates the effect of changing from only the importing (exporting) country having a membership to joint membership ($Bothin = 1$). For brevity, we report only the estimation results that control for all of TFEs, CPFES and MRTs. We find positive, economically and statistically significant effects on trade between country-pairs regardless of whether new members are importing or exporting countries in the first stage. In fact, the difference between the new member being an importer or an exporter is statistically indiscernible. The positive WTO trade effects for an exporting country in a single membership country-pair is plausible for a number of reasons: (i) if trade liberalization, especially in developing countries, is accompanied by administrative reform, the fixed costs of exporting can be reduced inducing a greater outflow of goods; (ii) as shown in Melitz (2003), only the most productive firms in the exporting country survive after trade liberalization, and thus the country's exports may increase with any trading partner; and (iii) some firms in exporting countries may upgrade technol-

ogy and improve their productivity with trade liberalization (see e.g. Lileeva and Trefler, 2010; Bustos, 2011; Aw, Roberts, and Xu, 2011) which may increase their comparative advantage even in respect to non-member importing countries.

We further distinguish between developing and industrial countries in Table 8. Subramanian and Wei (2007) suggest that the large exemptions for developing countries, especially before the Uruguay Round, would result in little impact of WTO membership on the trade of these countries. They indeed find strong WTO effects only for industrialized countries, which is consistent with their argument.

In the log-linear model with no zero observations, we obtain results that are qualitatively similar results to Subramanian and Wei (2007). However, in all other specifications, we find significant and strong WTO effects on trade in each stage for both developing and industrial countries. Especially, in the specification of the CPPML method with zero observations, the coefficients on WTO membership for developing and industrial countries are not statistically different.

4.3 WTO trade effects across decades

We also examine potential heterogeneity in WTO trade effects by doing the estimation decade by decade. Subramanian and Wei (2007) and Tang and Wei (2009) argue that the more substantial requirements of trade liberalization imposed by the WTO after its founding, especially in developing countries, would lead to more significant WTO effects in the 1990s than in the previous periods.

Table 9 show the results for each decade. Unlike Rose (2004), Eicher and Henn (2011), and Roy (2011), we find positive and statistically significant WTO effects in every decade in the all specifications except for the 1970s when the Tokyo Round (1973-1979) was concluded. The Tokyo Round was the biggest round up to that time with 102 participants and it covers the rules for non-tariff measures as well. However, as only seven (developing) countries (Surinam, Bangladesh, Hungary, Philippine, Romania, Singapore and Zaire) joined the WTO during this period, after controlling for CPFES the time variations from these countries may not be enough to identify WTO trade effects. In the specification of the CPPML method with zero observation, we find the largest overall WTO trade effects

occurring in the 1990s in line with Subramanian and Wei (2007), but only the effect of joint WTO membership is statistically significant. In other periods, WTO trade effects arise in the first-stage, but not in the second-stage. This may reflect the fact that under the old GATT rules, accession to the GATT required very few reforms for developing countries and many of them retained very high tariff rates even after becoming a GATT member (see Subramanian and Wei, 2007; Tang and Wei, 2009). The finding that positive WTO effects are evident only in the first stage suggests that focusing solely on joint WTO membership in investigating WTO trade effects, which is common in the literature, may be misguided.

5 Conclusions

This paper reexamines the effect of WTO membership on bilateral trade flows. First, we point out some previously unnoticed identification problems in the estimation of WTO trade effects. We show that controlling for MRTs, which is considered one of the most important steps to avoid omitted variable bias (OV) in the gravity model, would create a severe multicollinearity problem because of the structural properties of the two WTO membership variables. Results indicate near-perfect multicollinearity makes WTO trade effects estimates very fragile. A closer look at the data also reveals that a country-pairs' WTO membership status tends to change one membership at a time. This implies that comparing the WTO trade effects of joint memberships to that of no membership, as usually done in the literature, may lead to using little variations in identifying joint membership effects when CPFEs are controlled for. We propose a two-stage method that circumvents the tradeoff of OV with multicollinearity and the small number of observation problem. Once these identification problems are removed using the two-stage method, WTO trade effects are found to be consistently positive, and statistically and economically significant in all specifications. In several extensions of our specifications, we obtain robustly positive WTO effects. In particular, we find that both importing and exporting countries as new members enjoy positive WTO effects and so do both developing and industrial countries, which is in contrast to Subramanian and Wei (2007) and Eicher and Henn (2011).

Our findings suggest that the conflicting results of WTO trade effects in the previous literature could be attributable not only to OVB, as highlighted in Eicher and Henn (2011), or the model misspecification addressed in Chang and Lee (2011), but also to the multicollinearity of the two WTO membership variables as well as the identification problems with small observations.

The results of decade by decade estimations show the positive WTO effects in every decade except for the 1970s where there were small variations in the WTO memberships. It is also shown that except for the 1990s, WTO trade effects arise in the first-stage, not in the second-stage, implying that focusing only on joint WTO membership may understate WTO trade effects.

Table 1: Previous studies on WTO trade effects

	Rose(2004)	Tomz et al.(2007)	Subramanian & Wei(2007)	Liu(2009)	Eicher & Henn(2011)	Roy(2011)	Felbermayr & Kohler(2010)	Dutt et al.(2011)	Chang & Lee (2011)
Definition	<i>de jure</i>	<i>de facto</i>	<i>de jure</i> , mutually exclusive to PTAs and GSP	<i>de jure</i>	<i>de jure & de facto</i>	<i>de jure & de facto</i>	<i>de jure & de facto</i>	<i>de jure</i>	<i>de jure & de facto</i>
Other Classifications			Heterogeneity of WTO: Industrial or developing, before and after WTO for developing countries, Different sectors		Heterogeneous effects of individual PTAs		Extensive margin at country level	Intensive and Extensive margin at sectoral level	
Simultaneous inclusion of <i>Bothin & Onein</i>	Yes	Yes	No	Yes	No	Yes	No	No	.
Inclusion of zero observations	No	No	No	Yes	No	Yes	Yes	Yes	No
Primal Method	OLS	OLS	OLS	CPPML	OLS	OLS	PPML	OLS & Helpman et al.'s (2008) method	Matching methods (non-parametric)
Country-Time Fixed Effects or MRTs	No	No	CTFEs	Remoteness	CTFEs	CTFEs	CTFEs	CTFEs	MRTs (BB method)
Country-Pair Fixed Effects	Yes	Yes	No	Yes	Yes	No	No	Yes	.
Findings	little evidence	positive effects	strong effects for industrial, but not for developing countries	positive effects	no effects for either industrial or developing countries	no effects	1948-95: little evidence on intensive margins, but successful on extensive margins. 1995: positive effects	No effects on intensive margins but positive effects on extensive margins	positive effects

Table 2: WTO trade effects					
		Log-linear		(C)PPML	
		No zeros	With zeros	No zeros	With zeros
Controls					
TFEs	<i>Onein</i>	-0.20***	0.93***	-0.26***	-0.05
		(0.04)	(0.05)	(0.11)	(0.10)
	<i>Bothin</i>	-0.01	2.04***	-0.37***	-0.11
		(0.04)	(0.06)	(0.12)	(0.11)
TFEs, CPFEs	<i>Onein</i>	-0.06**	0.41***	0.10	NA
		(0.03)	(0.05)	(0.07)	
	<i>Bothin</i>	0.03	1.44***	0.30***	NA
		(0.03)	(0.06)	(0.08)	
TFEs, MRTs	<i>Onein</i>	2.03***	10.74***	0.33	0.88
		(0.23)	(0.43)	(0.58)	(0.58)
	<i>Bothin</i>	4.33***	20.84***	0.73	1.73
		(0.46)	(0.87)	(1.16)	(1.18)
TFEs, CPFEs, MRTs	<i>Onein</i>	0.68***	4.36***	NA	0.44
		(0.23)	(0.47)		(0.99)
	<i>Bothin</i>	1.46***	8.84***	NA	1.03
		(0.47)	(0.95)		(2.01)
	<i>RTA</i>	0.38***	0.53***	NA	0.39***
		(0.03)	(0.08)		(0.06)
No. of obs.		548,174	1,155,274	548,174	1,155,274

Notes: TFEs, CPFES, and MRTs denote time fixed effects, country-pair fixed effects and multilateral resistance terms based on the Baier and Bergstrand Method, respectively. Cluster (country-pairs) robust standard errors are reported in parentheses. ***, **, and * indicate that the coefficient is statistically significant at 1%, 5% and 10% levels, respectively. NA indicates that we are not able to obtain the value as the estimation does not converge. RTA represents regional trade agreements which include free trade agreements and customs union. When MRTs are included, the variables are "deMRTed" when entering the regression as described in equations (3) and (4). The difference in the number of observations used for the estimations between the log-linear model and the Poisson model arise because observations are dropped for the Poisson model estimation when there is only one observation for a pair and when dependent variables are all zero outcomes for a pair.

Table 3: The correlation between *Onein* and *Bothin*

	Whole sample	1950-60	60-70	70-80	80-90	90-2000
Unconditional	-0.58***	-0.21***	-0.46***	-0.66***	-0.70***	-0.82***
Conditional on TFEs	-0.64***	-0.19***	-0.46***	-0.58***	-0.61***	-0.73***
Conditional on TFEs and CPFES	-0.58***	-0.46***	-0.54***	-0.72***	-0.75***	-0.82***
Conditional on TFEs and MRTs	-0.99***	-0.99***	-0.99***	-0.99***	-0.99***	-0.99***
(Conditional on CTFEs)	NA	-0.99***	-0.99***	-0.99***	-0.99***	-0.99***
Conditional on TFEs, CPFES and MRTs	-0.99***	-0.99***	-0.99***	-0.99***	-0.99***	-0.99***
(Conditional on CPFES and CTFEs)	NA	-0.99***	-0.99***	-0.99***	-0.99***	-0.99***

Note: *** indicates the correlation is statistically significant different from zero at 1% level, respectively.

Table 4: The distribution of WTO membership amongst all country-pairs

Initial year	None	One	Both	None	None	None
Middle year					One	
Final year	None	One	Both	One	Both	Both
(%)	6.3	18.7	4.3	22.1	47	1.6

Table 5: Stage-by-stage WTO effects with formal WTO definition

Controls	(Reference group)		Log-linear		(C)PPML	
			No zeros	With zeros	No zeros	With zeros
TFEs	<i>(Nonein)</i>	<i>Onein</i>	-0.08**	1.00***	-0.20*	-0.04
			(0.03)	(0.05)	(0.12)	(0.11)
			[306,571]	[800,184]	[306,571]	[800,184]
	<i>(Onein)</i>	<i>Bothin</i>	0.16***	1.02***	-0.11*	-0.05
			(0.02)	(0.05)	(0.07)	(0.07)
			[492,328]	[906,842]	[492,328]	[906,842]
TFEs, CPFES	<i>(Nonein)</i>	<i>Onein</i>	-0.02	0.34***	0.13	0.19**
			(0.03)	(0.05)	(0.09)	(0.08)
			[306,571]	[800,184]	[302,807]	[799,232]
	<i>(Onein)</i>	<i>Bothin</i>	0.08***	0.81***	0.19***	0.21***
			(0.02)	(0.05)	(0.06)	(0.06)
			[492,328]	[906,842]	[489,578]	[906,316]
TFEs, MRTs	<i>(Nonein)</i>	<i>Onein</i>	0.01	0.76***	-0.06	0.07
			(0.02)	(0.03)	(0.07)	(0.07)
			[306,571]	[800,184]	[306,571]	[800,184]
	<i>(Onein)</i>	<i>Bothin</i>	1.05***	5.93***	0.24	0.18
			(0.08)	(0.21)	(0.21)	(0.25)
			[492,328]	[906,842]	[492,328]	[906,842]
TFEs, CPFES, MRTs	<i>(Nonein)</i>	<i>Onein</i>	-0.01	0.27***	0.11*	0.15***
			(0.02)	(0.03)	(0.06)	(0.05)
			[306,571]	[800,184]	[302,807]	[799,232]
		<i>RTA*</i>	0.26***	0.46***	0.19***	0.25***
			(0.05)	(0.11)	(0.09)	(0.10)
			[306,571]	[800,184]	[302,807]	[799,232]
	<i>(Onein)</i>	<i>Bothin</i>	0.30***	4.04***	0.87***	0.96***
			(0.06)	(0.19)	(0.22)	(0.21)
			[492,328]	[906,842]	[489,578]	[906,316]
		<i>RTA*</i>	0.36***	0.21***	0.39***	0.39***
			(0.03)	(0.08)	(0.05)	(0.05)
			[492,328]	[906,842]	[489,578]	[906,316]

Notes: See Table 1. Cluster (country-pairs) robust standard errors are reported in parentheses and the sample size are reported in square brackets.

Table 6: Stage-by-stage WTO effects with the WTO definition including nonmember participants

Controls	(Reference group)		Log-linear		(C)PPML	
			No zeros	With zeros	No zeros	With zeros
TFEs	<i>(Nonein)</i>	<i>Onein</i>	-0.19***	0.59***	-0.22*	-0.05
			(0.04)	(0.06)	(0.13)	(0.12)
			[222,939]	[568,482]	[222,939]	[568,482]
	<i>(Onein)</i>	<i>Bothin</i>	0.30***	0.63***	-0.09	-0.04
			(0.02)	(0.04)	(0.07)	(0.07)
			[519,287]	[1,068,778]	[519,287]	[1,068,778]
TFEs, CPFES	<i>(Nonein)</i>	<i>Onein</i>	0.03	0.63***	0.01	0.11
			(0.03)	(0.06)	(0.07)	(0.08)
			[222,939]	[568,482]	[219,956]	[567,437]
	<i>(Onein)</i>	<i>Bothin</i>	0.28***	0.36***	0.24***	0.23***
			(0.02)	(0.04)	(0.08)	(0.07)
			[519,287]	[1,068,778]	[516,230]	[1,068,226]
TFEs, MRTs	<i>(Nonein)</i>	<i>Onein</i>	0.01	0.59***	-0.08	0.05
			(0.03)	(0.05)	(0.08)	(0.07)
			[222,939]	[568,482]	[222,939]	[505,451]
	<i>(Onein)</i>	<i>Bothin</i>	1.58***	3.40***	0.16	0.34
			(0.09)	(0.18)	(0.28)	(0.28)
			[519,287]	[1,068,778]	[519,287]	[1,068,778]
TFEs, CPFES, MRTs	<i>(Nonein)</i>	<i>Onein</i>	0.05	0.55***	NA	0.10*
			(0.02)	(0.05)		(0.06)
			[222,939]	[568,482]		[567,437]
	<i>(Onein)</i>	<i>Bothin</i>	1.07***	0.92***	NA	0.97***
			(0.07)	(0.18)		(0.27)
			[519,287]	[1,068,778]		[1,068,226]

Notes: See Table 4.

Table 7: Stage-by-stage WTO effects with separate importer and exporter memberships

New member	(Reference group)		Log-linear		CPPML	
			No zeros	With zeros	No zeros	With zeros
Importer	<i>(None in)</i>	<i>Importer in</i>	0.02	0.33***	0.17**	0.22***
			(0.03)	(0.04)	(0.08)	(0.06)
			[173,214]	[523,617]	[170,330]	[522,400]
	<i>(Exporter in)</i>	<i>Both in</i>	0.25***	3.83***	0.74**	0.86***
			(0.08)	(0.24)	(0.31)	(0.28)
			[374,960]	[631,657]	[372,287]	[631,126]
Exporter	<i>(None in)</i>	<i>Exporter in</i>	-0.01	0.42***	0.11	0.16**
			(0.02)	(0.05)	(0.08)	(0.07)
			[189,203]	[524,999]	[186,290]	[523,787]
	<i>(Importer in)</i>	<i>Both in</i>	0.34***	4.49***	1.02***	1.13***
			(0.09)	(0.25)	(0.03)	(0.28)
			[358,971]	[630,275]	[356,809]	[629,710]

Notes: See Table 4. All regressions control for TFEs, CPFES and MRTs.

Table 8: Stage-by-stage WTO effects with separate importer and exporter memberships, developing (DEV) and industrial countries (IND)

New member	(Reference group)		Log-linear		CPPML	
			No zeros	With zeros	No zeros	With zeros
Importer	(None in)	<i>Importer in</i> DEV	-0.10	0.22***	0.18**	NA
			(0.03)	(0.05)	(0.07)	
			[173,214]	[523,617]	[170,330]	
		<i>Importer in</i> IND	0.16**	1.03***	0.16*	NA
			(0.07)	(0.15)	(0.08)	
			[173,214]	[523,617]	[170,330]	
	(Exporter in)	<i>Both in</i> DEV	0.10	1.78***	0.82***	1.10***
			(0.08)	(0.29)	(0.32)	(0.37)
			[374,960]	[630,257]	[356,809]	[629,710]
		<i>Both in</i> IND	0.37***	8.79***	1.01***	1.09***
			(0.09)	(0.41)	(0.35)	(0.28)
			[374,960]	[630,257]	[356,809]	[629,710]
Exporter	(None in)	<i>Exporter in</i> DEV	-0.05*	0.34***	0.09	0.14*
			(0.03)	(0.05)	(0.09)	(0.08)
			[189,203]	[524,999]	[186,290]	[523,787]
		<i>Exporter in</i> IND	0.22***	0.33***	0.21**	0.27***
			(0.05)	(0.04)	(0.07)	(0.08)
			[189,203]	[524,999]	[186,290]	[523,787]
	(Importer in)	<i>Both in</i> DEV	0.07	1.87***	0.71**	0.99***
			(0.15)	(0.29)	(0.25)	(0.28)
			[358,971]	[631,657]	[372,287]	[631,126]
		<i>Both in</i> IND	0.60***	7.00***	0.76**	0.86***
			(0.12)	(0.40)	(0.35)	(0.33)
			[358,971]	[631,657]	[372,287]	[631,126]

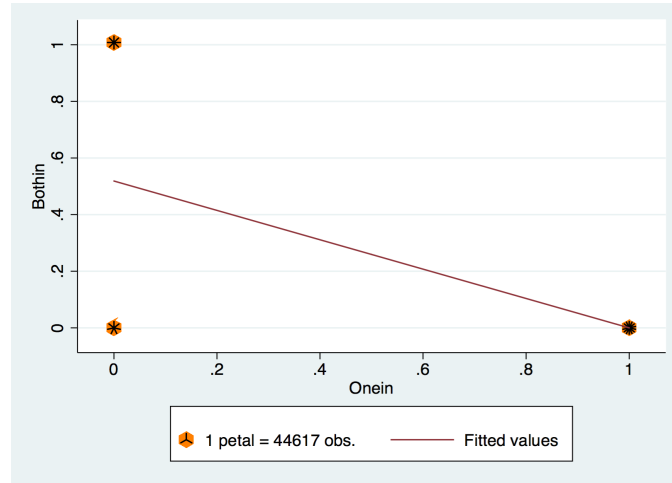
Notes: See Table 4. All regressions control for TFEs, CPFEs and MRTs.

Table 9: Across time WTO effects with formal WTO definition:with TFEs, CPFES and MRTs

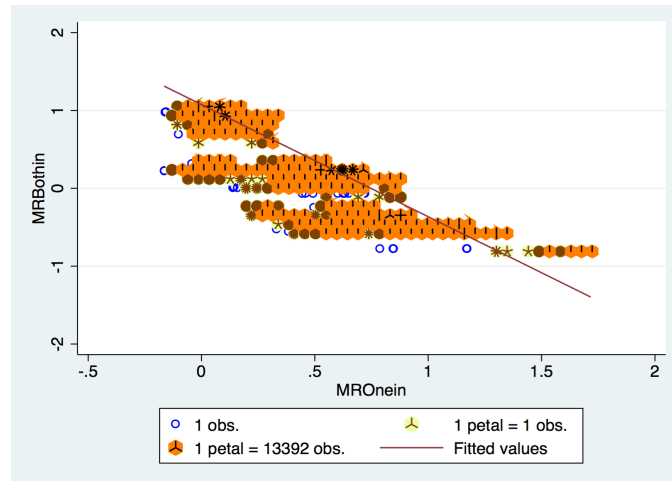
			Log-linear		CPPML	
(Reference group)			No zeros	With zeros	No zeros	With zeros
1948-1960	<i>(Nonein)</i>	<i>Onein</i>	0.12***	0.38***	0.11**	0.23***
			(0.03)	(0.06)	(0.12)	(0.05)
			[38,559]	[171,073]	[37,328]	[169,711]
	<i>(Onein)</i>	<i>Bothin</i>	0.21***	0.08	0.11	0.22
			(0.09)	(0.34)	(0.14)	(0.14)
			[37,647]	[93,080]	[36,951]	[92,327]
1960-1970	<i>(Nonein)</i>	<i>Onein</i>	-0.05**	0.09**	0.06***	0.15**
			(0.02)	(0.04)	(0.09)	(0.02)
			[47,964]	[161,177]	[45,785]	[157,845]
	<i>(Onein)</i>	<i>Bothin</i>	0.24***	2.58***	0.02	0.13
			(0.06)	(0.25)	(0.09)	(0.09)
			[61,162]	[143,581]	[59,430]	[141,350]
1970-1980	<i>(Nonein)</i>	<i>Onein</i>	0.09**	0.66***	-0.14***	NA
			(0.04)	(0.10)	(0.05)	
			[70,968]	[171,592]	[68,687]	
	<i>(Onein)</i>	<i>Bothin</i>	0.20	4.46***	-0.34**	-0.17
			(0.18)	(0.59)	(0.17)	(0.17)
			[103,048]	[197,077]	[101,251]	[195,720]
1980-1990	<i>(Nonein)</i>	<i>Onein</i>	0.05	0.43***	0.21***	0.21***
			(0.03)	(0.07)	(0.06)	(0.06)
			[79,341]	[108,854]	[76,909]	[180,177]
	<i>(Onein)</i>	<i>Bothin</i>	0.45***	3.08***	0.19	0.18
			(0.13)	(0.33)	(0.12)	(0.21)
			[125,300]	[224,035]	[122,993]	[223,129]
1990-2001	<i>(Nonein)</i>	<i>Onein</i>	-0.04	0.18***	0.02	0.01
			(0.03)	(0.05)	(0.05)	(0.05)
			[94,306]	[180,327]	[90,736]	[178,792]
	<i>(Onein)</i>	<i>Bothin</i>	0.19*	1.35***	0.53***	0.56***
			(0.10)	(0.21)	(0.13)	(0.13)
			[200,850]	[316,321]	[197,974]	[315,762]

Notes: See Table 4. All regressions control for TFEs, CPFES and MRTs.

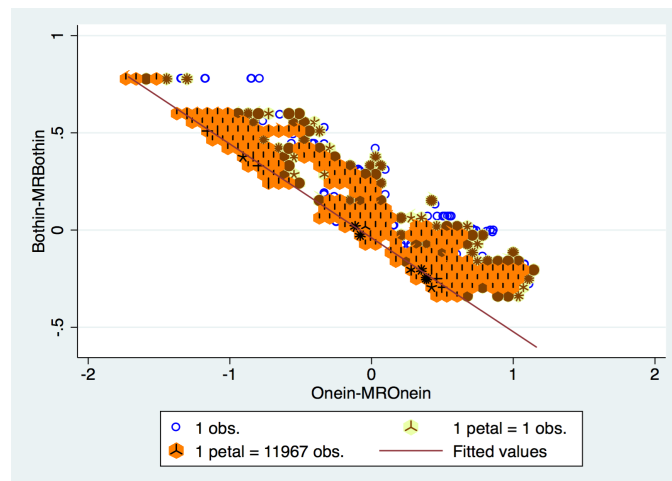
Figure 1: The correlation between *Onein* and *Bothin*



(a) *Onein* and *Bothin*



(b) *MROnein* and *MRBothin*



(c) *Onein - MROnein* and *Bothin - MRBothin*

Appendix

This appendix provides an explanation of why the two WTO membership variables, *Onein* and *Bothin* are of near-perfect multicollinearity after controlling for MRTs using the BB method. Recall that the BB method involves transforming z into $(z - MRz)$, where $z_{ijt} \in \mathbf{Z}_{ijt} \equiv (\mathbf{X}_{ijt}, \text{Onein}_{ijt}, \text{Bothin}_{ijt})$ and MRz is a linearized MRT given by (4):

$$MRz_{ijt} = \sum_{k=1}^N \left(\frac{Y_{kt}}{Y_{Wt}} \right) z_{ikt} + \sum_{m=1}^N \left(\frac{Y_{mt}}{Y_{Wt}} \right) z_{jmt} - \sum_{k=1}^N \sum_{m=1}^N \left(\frac{Y_{kt}}{Y_{Wt}} \right) \left(\frac{Y_{mt}}{Y_{Wt}} \right) z_{kmt}$$

Also recall that for any dyad observation, there are only three possible membership statuses: no membership ($\text{Onein} = 0, \text{Bothin} = 0$), a single membership ($\text{Onein} = 1, \text{Bothin} = 0$), or joint memberships ($\text{Onein} = 0, \text{Bothin} = 1$). Table 10 shows the values for $MR\text{Onein}$, $MR\text{Bothin}$, $(\text{Onein} - MR\text{Onein})$ and $(\text{Bothin} - MR\text{Bothin})$ for each of these three cases.¹³

For the no membership case, each of the first two components of $MR\text{Onein}$ is equal to the sum of the GDP shares of all member countries at time t and we denote this sum as M_t . Likewise, for the joint membership case, each of the first two components of $MR\text{Onein}$ is equal to the sum of the GDP shares of all non-member countries at time t and we denote this sum as N_t . For the single membership case, the sum of the first two terms of $MR\text{Onein}$ is equal to 1 as they are the sum of the GDP shares of all member and non-member countries, i.e. $M_t + N_t = 1$. The last component of $MR\text{Onein}$, however, is the same for all three cases because it represents the (weighted) membership statuses of all country-pairs at time t . The values for $MR\text{Bothin}$ for each of the three cases can be explained in a similar way.

Now consider panel (c) of Figure 1 where $(\text{Bothin} - MR\text{Bothin})$ is plotted against $(\text{Onein} - MR\text{Onein})$ using data for multiple periods. To keep the calculation simple, we assume $Y_{it} = Y_{jt}$ and each country is small enough so that $\frac{Y_{it}}{Y_{Wt}}$ is negligible. Then, for a given time t there can only be three possible

¹³We do not include Onein_{ii} and Bothin_{ii} in the calculation as they are not definable.

Table 10: The values of the two membership variables

WTO Membership	<i>Onein</i>	<i>Bothin</i>	<i>MROnein</i>	<i>MRBothin</i>	<i>Onein</i> – <i>MROnein</i>	<i>Bothin</i> – <i>MRBothin</i>
No membership	0	0	$2M_t - a_t$	$-b_t$	$-2M_t + a_t$	b_t
Single membership	1	0	$1 - a_t$	$M_t - b_t - \frac{Y_{it}(or Y_{jt})}{Y_{Wt}}$	a_t	$-M_t + b_t + \frac{Y_{it}(or Y_{jt})}{Y_{Wt}}$
Joint memberships	0	1	$2N_t - a_t$	$2M_t - b_t - \frac{Y_{it} + Y_{jt}}{Y_{Wt}}$	$-2N_t + a_t$	$1 - 2M_t + b_t + \frac{Y_{it} + Y_{jt}}{Y_{Wt}}$

Note: M_t and N_t denote the sum of the GDP shares of all the member countries and all the non-member countries in the world at time t , respectively. a_t and b_t denote the last term of *MROnein* and *MRBothin* at time t , respectively.

values for the two variables, i.e. there will only be three points on the scatter plot for time t . Let us denote these points as A (for no membership), B (for a single membership), and C (for joint memberships). From Table 10, the slope between A and B is $-\frac{M_t}{2M_t} = -0.5$, and the slope between B and C is $-\frac{1-M_t}{2N_t} = -0.5$ because $M_t + N_t = 1$. This means that the three points lie on a downward sloping straight line with a slope of -0.5 and, thus, there is a prefect negative cross-sectional correlation between $(Bothin - MRBothin)$ and $(Onein - MROnein)$. Given that for the majority of countries their world GDP share weights and their variations across countries are small, the cross-sectional correlation between the two variables must be very close to -1.

Over time, because of changes in membership statuses or GDP shares, M_t , N_t , x_t and y_t change and so do *MROnein* and *MRBothin*. These changes will affect only the intercept of the line on which points A, B and C lie, but not the slope. However, because *MROnein* and *MRBothin* are dominated by large countries' membership statuses, due to the GDP share weights, the fact that most large countries have been WTO members from the very beginning, and that the GDP shares of most countries change only very slowly, the changes in the intercept over-time are small.

The panel (c) of Figure 1 clearly shows these findings. The slope of the fitted value line is very close to -0.5 and the majority of the observations are distributed in a direction that runs parallel to this line, causing near-perfect multicollinearity.

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