

# Does the WTO Promote Trade? A Meta-analysis and New Structural Gravity Estimates\*

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## Abstract

The World Trade Organization (WTO) and its predecessor, the General Agreement on Tariffs and Trade (GATT), are key institutions of the multilateral trading system. While the WTO is expected to promote trade by reducing tariffs and non-tariff barriers, existing estimates of its effect on trade vary widely in magnitude, sign, and significance. We collect 2,547 estimates from 71 papers and apply meta-analysis techniques to conduct a systematic quantitative review of the literature, complementing it with established advances in gravity models to obtain estimates of the WTO's impact on trade. The meta-analysis shows that, on average, the literature finds a significant and positive trade effect of the WTO, although the estimates depend strongly on study characteristics. Moreover, we find no evidence of publication bias. Our structural gravity estimates confirm these findings: the WTO increases trade. However, the effects are heterogeneous across sectors and income levels of trading partners.

**Keywords:** World Trade Organization, trade, gravity model, meta-analysis.

**JEL codes:** C83, F13, F14, F15.

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\*An earlier version of this paper circulated under the title: “Star Wars - Return of the WTO trade effect”. The title referred to the debate in the literature about the statistical significance of the coefficient, as usual indicated by one, two or three “stars” after the number. It also circulated under the title: “Does the WTO Promote Trade? A Meta-analysis”. We thank Pablo Millan Perez and Miguel Ibañez Puertas for excellent research assistance and Raquel Lorenzo for her assistance in thoroughly double-checking the data used in the meta-analysis. In writing this paper we have benefited from comments by Tomas Havránek, Roberta Piermartini, and by participants at the WTO Public Forum, XXV Conference on International Economics and XII Meeting on International Economics, INTECO, and the Banco de España research seminar. The views expressed in this paper are those of the authors and do therefore not necessarily reflect those of the Banco de España or the Eurosystem. Conflict of interest: none.

# 1 Introduction

January 1, 2025 marked the 30th anniversary of the creation of the World Trade Organization (WTO), the successor to the General Agreement on Tariffs and Trade (GATT) signed almost 50 years earlier in 1947. These institutions are responsible for ensuring that their members have common rules for trade in goods and services, both in terms of tariffs and non-tariff measures, and for creating fora to resolve trade disputes, discuss new trade rules, and promote transparency in trade policy matters.

Yet, despite its long institutional history and the array of policies it has introduced, the extent to which the GATT/WTO multilateral trading system promotes trade remains a subject of ongoing debate.<sup>1</sup> Indeed, since Rose (2004) first estimated a limited or null role for the WTO in promoting trade, a plethora of researchers have debated the validity of his findings, and whether they depend on the details of the methodology used. For example, the literature has argued that the size of the WTO trade effect depends on: the characteristics of the trading partners (Subramanian and Wei, 2007); whether trade policy changes other than those directly linked to WTO membership, e.g. trade agreements, are taken into account (Tomz et al., 2007); the set of fixed effects used in the estimation strategy (Roy, 2011; Cheong et al., 2014); or the use of both domestic and international trade data (Larch et al., 2025a).

In this context, this paper makes two contributions. First, we extend the meta-analysis literature in the field of international economics, by examining the WTO trade effect using meta-analysis techniques. Meta-analysis in economics has proven to be a successful tool for providing a quantitative summary of existing debates in the literature, often challenging previous findings or providing new insights into the size of the effect being analyzed (Gechert, 2022). Efforts in this direction have been made for the effects of national borders (Havránek and Irsova, 2017) and currency unions (Havránek, 2010) on trade, the effects of foreign direct investment on productivity (Havránek and Irsova, 2011; Iršová and Havránek, 2013; Hampl et al., 2020), or to revise the value of the elasticity of substitution between domestic and foreign goods, a key parameter in many trade models (Bajzik et al., 2020).<sup>2</sup> We collect 2,547 estimates of the WTO trade effect from 71 papers and construct several indicators that categorize the salient features of the research design employed for each of the estimates. Second, guided by theory and capitalizing on established developments in the empirical trade literature, we obtain an updated set of WTO effects

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<sup>1</sup>For the sake of simplicity, we will use “WTO” in the text for referring to GATT/WTO.

<sup>2</sup>Similarly, meta-analysis techniques have also been applied in the field of macroeconomics and macroeconomic policy. For example, Rusnak et al. (2013) and Havránek and Rusnak (2013) examine the impact of monetary policy on prices; Babecky and Havránek (2014) analyze the effects of structural reforms on growth; Havránek et al. (2015) look at the elasticity of inter-temporal substitution in consumption, a key parameter in many macroeconomic models.

and explore their heterogeneity across several dimensions.

The meta-analysis reveals three important insights. First, despite the doubts raised by Rose’s original paper, the cumulative literature shows that the effect of WTO membership on bilateral trade—the “WTO trade effect”—is significantly positive for trade between members and between members and non-members. The average size of these trade effects is about +23% and +14%, respectively. Second, the average estimate masks considerable heterogeneity in the estimates related to study characteristics. Indeed, when controlling for study characteristics, one of the most important being domestic trade, the “WTO trade effect” increases to approximately +50%, aligning closely with our structural gravity estimates. Third, the empirical literature estimating the WTO effect does not suffer from publication bias. Publication bias refers to the systematic distortion or omission of research results that are not considered important or statistically significant.<sup>3</sup> Publication bias typically occurs when studies with the expected sign or statistically significant results are more likely to be published or given prominence, and studies with unexpected or statistically insignificant results are not published. The absence of publication bias in the WTO effect may be a by-product of the fact that the first study found insignificant results, leading to a lively debate with several studies supporting findings in opposite directions.

Although the meta-analysis is a powerful tool for summarising existing estimates of the WTO’s trade effects, it is inherently shaped by the methods and data used in the original studies. Indeed, meta-analyses often provide a means to understand how different research practices shape results. In this paper, we take a step further by examining not only how following best practices affects the estimated WTO effect, but also by estimating modern, state-of-the-art, gravity models with several different databases. Thus, we complement our meta-analysis by providing updated estimates of the WTO trade effect based on a state-of-the-art gravity model. A key feature of this specification is the inclusion of domestic trade flows, which is consistent with the theoretical foundations of the gravity model (Yotov, 2022).

In doing so, we push the boundaries of the literature along several dimensions: first, we extend the time dimension of our analysis, in some cases starting from 1948. This is important because the closer we get to the present, the fewer countries that are not GATT or WTO members, the smaller the reference group, and the less variation remains from the fixed effects included in typical gravity models. Second, we go beyond the manufacturing sector, which has been the focus of some recent papers analyzing the WTO effect using state-of-the-art techniques (Felbermayr et al., 2024; Larch et al., 2025a). Indeed, we

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<sup>3</sup>Rothstein et al. (2005) define publication bias as the “tendency to decide to publish a study based on the results of the study rather than on its theoretical or methodological quality”.

include agriculture, mining, manufacturing, and services sectors in our analysis. Third, we compare two different specifications, which can be understood as upper- and lower-bound estimates of the WTO effect.<sup>4</sup> Fourth, we draw on nine different databases to obtain our estimates. Seven are used for the main analysis based on aggregate or manufacturing bilateral trade flows, while two additional databases are employed for sector-level results. This ensures that our findings on the WTO effect are not driven by any particular data source. Our estimates confirm the main findings of the meta-analysis: the WTO trade effect is large, positive, and significantly different from zero. Our preferred lower- and upper-bound specifications suggest that the increase in trade between WTO members is in the range of +35% to +51%. However, these effects are heterogeneous across sectors and income levels.

The WTO has a large, positive, and significant effect on trade in goods and trade in services, although the effect tends to be larger for the former. Indeed, the trade effect of the WTO in agriculture and manufacturing is about two to three times as large as its effect on services. The effects on trade between advanced economies (AEs) and emerging market and developing economies (EMDEs), or among EMDEs, tend to be larger than those on trade between AEs. Combining these two dimensions, our results suggest that in agriculture, WTO membership tends to favour more exports from EMDEs (both to AEs and to other EMDEs). In mining and energy, WTO membership tends to favour more exports from AEs to EMDEs. In manufacturing, WTO membership favours exports across the board, i.e., all combinations of trade by level of development (i.e., AEs-AEs, AEs-EMDEs, EMDEs-AEs, and EMDEs-EMDEs). The results for services are qualitatively similar to those for manufacturing but of a smaller magnitude.

The remainder of the paper is organized as follows: Section 2 offers a brief review of the empirical model, which is typically used to obtain estimates of the effects of GATT and WTO; Section 3 describes how we compiled the estimates available in the literature; Section 4 characterizes the distribution of available estimates and studies whether there is evidence of publication bias; Section 5 analyzes how various characteristics affect the magnitude of the WTO estimates. Finally, in Section 6, we estimate the impact of the WTO effect on trade using up-to-date practices for gravity estimation. Section 7 concludes.

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<sup>4</sup>The upper- and lower-bound specifications differ by the inclusion of time-varying dummy variables distinguishing international from domestic trade flows (see Section 6). Although the impact of these dummies on the estimates is uncertain, given the strong rise in international trade during most of our sample, for simplicity we refer to the two specifications as “upper bound” and “lower bound” throughout the paper.

## 2 Estimates of the WTO effect

Membership in the WTO often requires the implementation of substantial trade liberalization reforms (e.g., reduction of tariffs and non-tariff measures) and implies easier access to members' markets. Thus, the GATT and WTO are expected to reduce trade barriers, especially between members, and to promote a reciprocal opening of economies to trade. However, in a very influential paper, Rose (2004) reported estimates that challenged the view that WTO membership promotes trade among members. Since then, dozens of researchers have attempted to either refute or rationalize his findings.

Rose (2004), and most subsequent studies, use a gravity model to analyze this question. Gravity models—whose theoretical foundations were laid by Anderson (1979) and Bergstrand (1985, 1989), which have become the workhorse model of trade (Eaton and Kortum, 2002; Anderson and van Wincoop, 2003; Arkolakis et al., 2012)—explain trade flows between two locations in terms of their economic size, distance, and embedding into the world economy, the latter two being proxies for trade costs. Researchers using gravity models often estimate regressions of the form:

$$\log X_{ijt} = \gamma_1 \textit{Bothin}_{ijt} + \gamma_2 \textit{Onein}_{ijt} + \beta Z_{ijt} + \varepsilon_{ijt}, \quad (1)$$

where  $X_{ijt}$  denotes bilateral trade flows from exporter  $i$  to importer  $j$  over period  $t$ ,  $\textit{Bothin}$  is an indicator that countries  $i$  and  $j$  are both GATT or WTO members at time  $t$ ,  $\textit{Onein}$  is an indicator that only  $i$  or  $j$  is a member at time  $t$ ,  $Z_{ijt}$  is a vector of other characteristics included in the specification, and  $\varepsilon_{ijt}$  is a remainder error term.

Because Equation (1) is a log-linearized transformation of the equation implied by modern trade theory, many researchers, especially in more recent times, have relied on the Poisson Pseudo Maximum Likelihood (PPML) estimator (Santos Silva and Tenreyro, 2006) to estimate the following version of the same equation, where trade is measured in levels:

$$X_{ijt} = \exp(\gamma_1 \textit{Bothin}_{ijt} + \gamma_2 \textit{Onein}_{ijt} + \beta Z_{ijt}) + \varepsilon_{ijt}. \quad (2)$$

Regardless of whether the equation to be estimated is (1) or (2), the parameters of interest  $\gamma_1$  and  $\gamma_2$  have the same interpretation in both cases: they measure the semi-elasticity of bilateral trade to WTO membership.<sup>5</sup>

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<sup>5</sup>When the indicators  $\textit{Bothin}$  and  $\textit{Onein}$  are included simultaneously in the regression and data on domestic trade

### 3 Literature search and number of observations included in the meta-analysis

To start with a large sample of papers, we first identified all documents that cite Rose (2004) using Google Scholar. We performed this search on Google Scholar on March 30, 2022, using the software Publish or Perish, version 8. We reduced the universe of eligible papers in a step-wise fashion. In the first step, we eliminated books, dissertations, and papers not written in English. We also eliminated papers not accessible online. This left us with 1,821 candidate papers. In a second step, we screened the papers by reading the title, abstract, and introduction to determine whether they contained estimates of the effect of WTO membership on trade. 546 papers survived this screening step and were considered eligible. The flowchart in Figure 1 shows the information flow at each stage of the literature search, including the number of studies identified, screened, and deemed eligible.

We used three criteria to select estimates from the eligible papers. First, at least one estimate in the paper must be from a gravity equation that regresses trade on a dummy variable indicating WTO membership (either *Bothin* or *Onein*). Second, to be consistent with equations (1) or (2), the dependent variable must be either the level of trade flows or the log of trade flows, so that the parameter can be interpreted as a semi-elasticity. Hence, we excluded regressions using trade growth, trade volatility, or export and import ratios. Third, we excluded estimations that did not estimate the total effect of trade and focused only on the extensive or intensive margin.<sup>6</sup> If there were multiple versions of the paper, we used the estimates reported in the version published in a journal or, if unpublished, the most recent version.

Our final dataset contains 2,547 estimates from 71 papers. A total of 1,569 of these estimates correspond to the coefficient for the *Bothin* variable and 978 to the coefficient of the *Onein* variable. We record the estimated coefficient and the reported standard error when available. There are 2,284 cases with valid

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is not used, as in the original article by Rose, the comparison is relative to the group of country pairs in which neither country is a member of the WTO. In later studies, the indicator *Onein* tends to be omitted while *Bothin* is included. The reason is that later papers tend to include country-time fixed effects, a theory-consistent way of accounting for multilateral resistance terms (MRTs), which are multicollinear with WTO indicators. In these cases, the interpretation of *Bothin* is the same as before. If there are no controls to absorb the effect of *Onein*, then the parameter of *Bothin* estimates the effect on trade of both the exporter and the importer being members of the WTO relative to the counterfactual of at least one of them not being a member. This distinction is often not made by the authors of the papers, who refer to any of the estimates as “the WTO effect”. In line with this interpretation, we include all regressions in which at least one of *Bothin* or *Onein* is present.

<sup>6</sup>We excluded estimations when the extensive margin was defined as the number of firms that participate in trade, or products that are traded, and the intensive margin was defined as the ratio of trade value to the number of firms/products. This does not mean that we automatically excluded estimations based on trade flows being strictly positive, as occurs, for example, when the logarithm of trade value is used as the dependent variable.

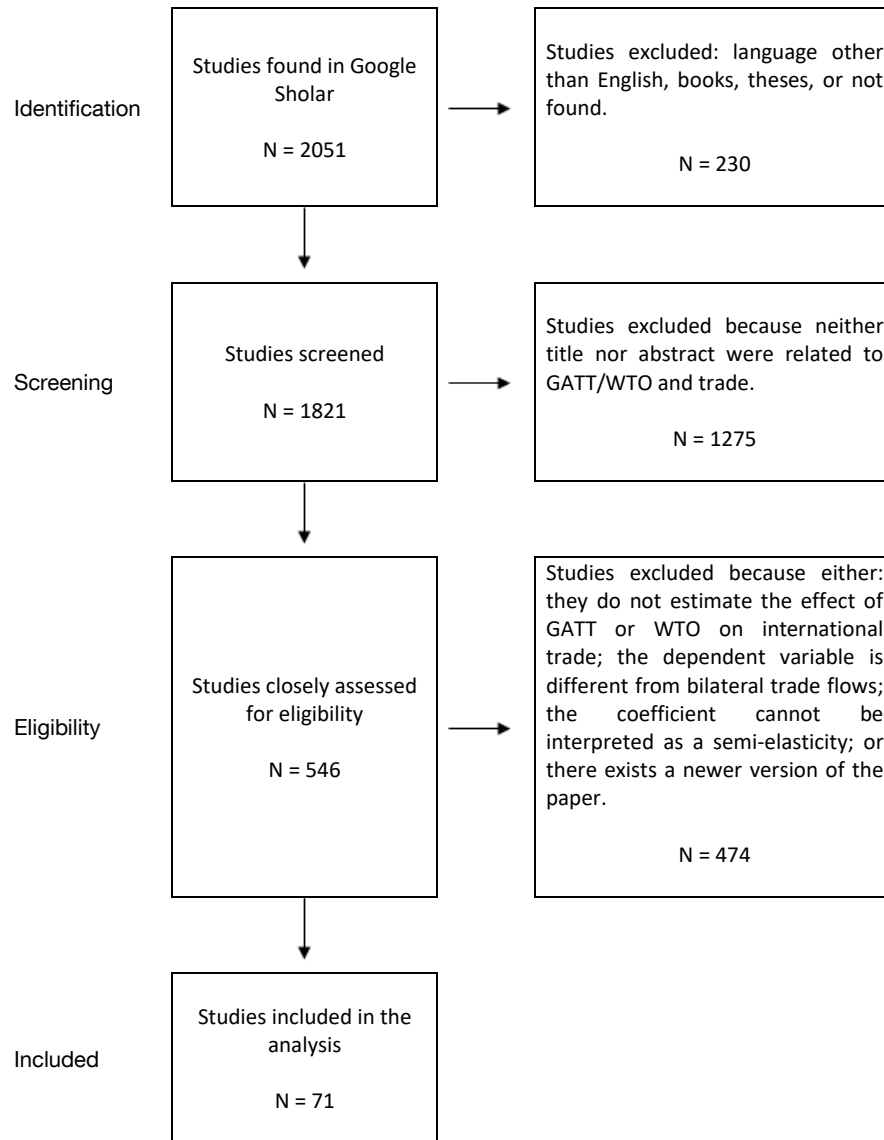


Figure 1: Identification, screening and eligibility of studies

non-missing observations for the standard error. Out of these, 1,422 cases (from 62 studies) correspond to the coefficient for the *Bothin* variable and 862 cases (from 34 studies) to the coefficient of the *Onein* variable. In addition to coefficients and standard errors, we collect information on several additional variables (described in a later section) that capture characteristics of the specification used, the estimation method, or information about the papers in which the regressions are included. Most of the coefficients in the final sample come from publications in peer-reviewed articles. They make up for 2,207 coefficients (1,399 for *Bothin* and 808 for *Onein*). Out of these, 2,188 have valid non-missing observations for the standard error (1,381 for *Bothin* and 807 for *Onein*).

## 4 Publication bias

In Table 1, we report unweighted and weighted means of the two coefficients of interest and confidence intervals as a measure of the variation in the coefficients. These confidence intervals are obtained by regressing the coefficient on a constant and using standard errors clustered by study to construct the asymptotic distribution of the coefficients around the mean.<sup>7</sup> The table shows that the literature finds point estimates of the effect of WTO membership on bilateral trade—the “WTO trade effect”—that are positive on average for trade between members and between members and non-members. Their magnitudes are centred around 0.21 and 0.13, respectively, implying an increase of trade associated with WTO membership of  $\exp(0.21) - 1 = 23\%$  and  $\exp(0.13) - 1 = 14\%$ . When each study is weighted equally, irrespective of how many estimates it provides, the resulting averages remain very similar.

Table 1: Average estimates in the literature

|                                           | Studies | Mean | 95% CI      | Weighted mean | 95% CI      |
|-------------------------------------------|---------|------|-------------|---------------|-------------|
| <i>Bothin</i> ( $\gamma_1$ ): Both in WTO | 65      | 0.21 | [0.14 0.28] | 0.20          | [0.15 0.26] |
| <i>Onein</i> ( $\gamma_2$ ): One in WTO   | 35      | 0.13 | [0.05 0.21] | 0.12          | [0.05 0.20] |
| <i>All</i>                                | 71      | 0.18 | [0.12 0.24] | 0.17          | [0.12 0.22] |

**Note:** The weighted means weight all studies equally, regardless of how many estimates they report. Confidence intervals are calculated by clustering standard errors at the study level.

While positive on average, the estimates of the “WTO trade effect” vary widely in the literature (Figure 2), with coefficient values ranging from  $-5.67$  to  $+8.01$ .<sup>8</sup> The histogram of the empirical distribution of

<sup>7</sup>To ensure that these results are not driven by small clusters, we calculate the intervals without study-level clustering of standard errors. See Table A.1 in the Appendix for more details.

<sup>8</sup>See Figure A.1 in the Appendix for a histogram of t-statistics.

the estimated coefficients has a peak near but above zero for both coefficients of interest (*Bothin* and *Onein*). In the case of *Bothin*, negative coefficients are very close to zero in most cases, while for *Onein* there are several cases where negative coefficients are far from zero. The differences in the empirical distributions of *Bothin* and *Onein* may reflect their different theoretical interpretation. For *Bothin*, there is little theoretical justification for a negative effect because WTO membership reduces bilateral trade costs by reducing trade policy barriers between trading partners, and is therefore expected to increase trade. The expected sign of *Onein*, on the other hand, can go in either direction, as WTO membership of only the exporter or the importer can, in principle, lead to trade diversion because, once a member, it is more convenient than before to trade with other members, and less convenient than before to trade with non-members.<sup>9</sup> It can also lead to trade creation because (at least part of) WTO members' trade reforms can be unilateral, thus opening the economy to members and non-members, or because of dynamic consequences of trade that make members' exports more competitive. Interestingly, estimates vary widely not only between studies but also within studies (see Figures A.2 to A.5 in the Appendix).

The literature suggests several approaches to detect publication bias. It is common to start with a graphical analysis (Egger et al., 1997) using “funnel plots”. This technique uses a scatter plot with the estimated size of an effect on the horizontal axis and the precision of the estimates (calculated as the inverse of the standard errors) on the vertical axis. Without publication bias, the scatter plot should lie within lines that look like an inverted funnel: estimates with high precision should be close to the mean or median effect, whereas those with low precision will be more dispersed. The distribution should also be symmetric. Figure 3 does not suggest the presence of publication bias, as the scatter plots for the effect of WTO membership on trade resemble inverted funnels and do not show the pattern typical of publication bias (high precision estimates far from the mean or median). Moreover, the funnel plots do not show an obvious difference between published (shown in black) and unpublished results (shown in blue).

A test of publication bias that goes beyond a visual assessment consists of regressing the estimated coefficients on the standard errors of the estimates. The idea behind this regression is that publication bias leads to a co-movement between the magnitude of estimates and standard errors, which can be detected by finding a non-zero coefficient associated with the standard errors (Stanley, 2005; Havránek,

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<sup>9</sup>This, of course, is related to the use of country-time fixed effects. If the latter are included, they control for the traditional general equilibrium trade diversion effects. Thus, we would expect that the chance of obtaining a positive estimate on *Onein* is higher when including country-time fixed effects, which is only possible when additionally domestic trade is included in the dataset.

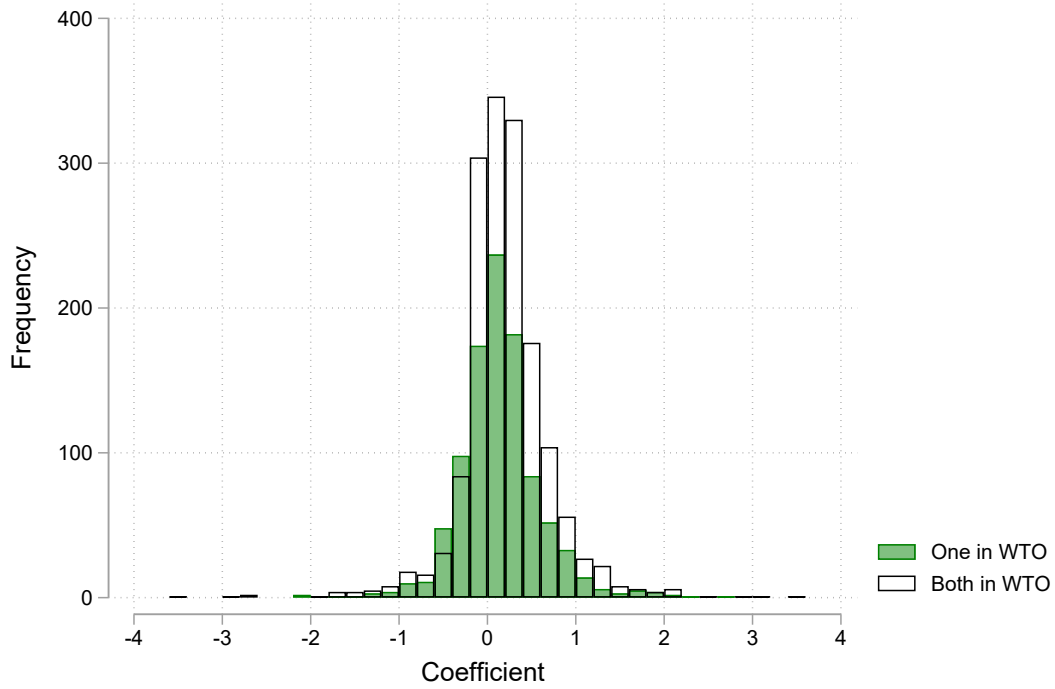


Figure 2: Histogram of coefficients

**Note:** The figure shows the histogram of the estimates of the WTO effect on trade reported in individual studies.

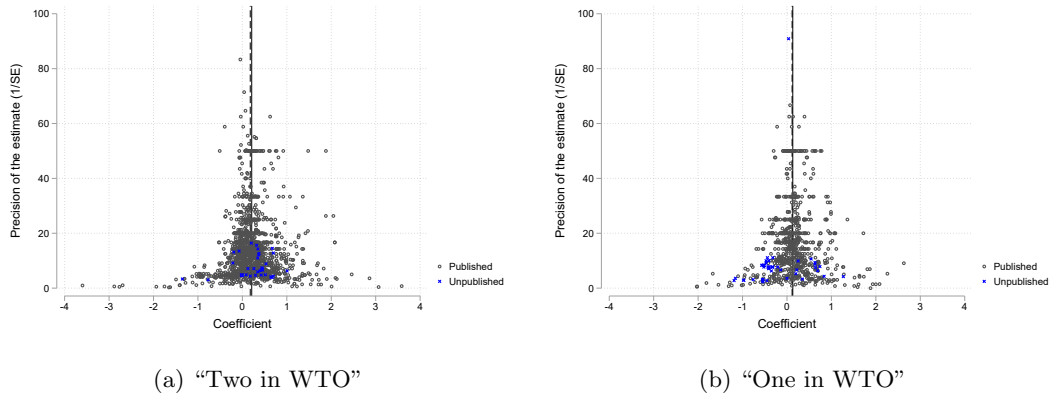


Figure 3: Funnel plots

**Note:** The figures show scatter plots with the estimated coefficient on the horizontal axis and precision (the inverse of the standard error) on the vertical axis. Estimates that belong to published articles are shown in black and those that belong to unpublished work are shown in blue. The figure on the left shows estimates for *Bothin* and the figure on the right for *Onein*.

2010; Havránek and Irsova, 2011; Bajzik et al., 2020). More specifically, the method suggests a regression specified as follows:

$$WTO\ effect_{ij} = \alpha + \beta \times \mathbf{se}(WTO\ effect)_{ij} + \epsilon_{ij}, \quad (3)$$

where the left-hand side variable,  $WTO\ effect_{ij}$ , identifies the estimate  $i$  of the WTO trade effect in study  $j$ . On the right-hand side of the equation, the explanatory variables are a constant with coefficient  $\alpha$ , which captures the mean effect excluding publication bias, and  $\mathbf{se}(WTO\ effect)_{ij}$ , the reported standard error of the WTO effect, whose coefficient  $\beta$  identifies the size and sign of the publication bias. Intuitively, if  $\beta = 0$ , there is no evidence of publication bias. If  $\beta \neq 0$  instead, it means that the estimates and their standard errors are correlated, a finding that suggests the presence of publication bias.

Table 2 shows the results of estimating Equation (3) using different methods. Columns (1) to (3) use the sample of *Bothin* coefficients. Column (1) reports the results of a pooled set of coefficients, including all estimates of the type *Bothin*. Column (2) restricts the sample to estimates published in a journal. Column (3) includes fixed effects by study. Columns (4) to (6) mimic columns (1) to (3), but using the sample of *Onein* coefficients.

In the case of *Bothin* (columns (1) to (3)), the data suggest no publication bias, as there is no conclusive evidence that the standard errors are systematically related to the coefficient estimates. In the case of *Onein* (columns (4) to (6)), there is also no clear evidence of publication bias. Indeed, publication bias is typically identified when there is strong and consistent evidence, such as numerous regressions showing a significant correlation between standard errors and the size of the coefficient. This is not the case here: only in a few instances is the coefficient marginally significant. Furthermore, as shown in the Appendix, this marginal significance is driven by outliers and by the presence of small clusters, which disappear once the sample is winsorized at 1% or clustering at the study level is avoided.<sup>10</sup> The point estimates are in the range of 0.26–0.28 after controlling for standard errors.

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<sup>10</sup>In further robustness tests, reported in Tables A.3 and A.4 in the Appendix, we run the same regressions on a sample winsorized at the 1% level to ensure that results are not driven by outliers, and for different sub-samples of well-identified studies, i.e., those studies that use structural gravity models including all theoretically grounded fixed effects. We also show the same regressions for the rest of the sample, i.e., those studies that do not use structural gravity models. We also use alternative tests of publication bias that relax different assumptions. First, we use the estimator proposed by Ioannidis et al. (2017), which consists of a weighted average of adequately powered (WAAP) results. The use of this estimator is likely to reduce bias due to publication selection, reporting, or small sample bias. Second, we use the Andrews and Kasy (2019) selection model, which relaxes the assumption that the bias is a linear function of the standard error. Third, we use the meta-analysis instrumental variable estimator (MAIVE) proposed by Havránek et al. (2023), which relaxes the assumption that the standard error is exogenous to the research studies and therefore reports estimates that are robust to the standard error being affected by estimation techniques or other issues of the sort. Results are in line with those reported in the main text and are reported in Tables A.5, A.6, and A.7, respectively.

Table 2: Tests for publication bias

| VARIABLES      | (1)<br><i>Bothin</i><br>All | (2)<br><i>Bothin</i><br>Journals only | (3)<br><i>Bothin</i><br>FE | (4)<br><i>Onein</i><br>All | (5)<br><i>Onein</i><br>Journals only | (6)<br><i>Onein</i><br>FE |
|----------------|-----------------------------|---------------------------------------|----------------------------|----------------------------|--------------------------------------|---------------------------|
| Standard error | -0.314*<br>(0.189)          | -0.311<br>(0.196)                     | -0.436<br>(0.340)          | 0.002<br>(0.312)           | 0.002<br>(0.291)                     | 0.002<br>(0.352)          |
| Constant       | 0.258***<br>(0.041)         | 0.257***<br>(0.042)                   | 0.275***<br>(0.058)        | 0.138**<br>(0.058)         | 0.159***<br>(0.051)                  | 0.138**<br>(0.062)        |
| Observations   | 1,422                       | 1,381                                 | 1,422                      | 862                        | 807                                  | 862                       |
| R-squared      | 0.019                       | 0.019                                 | 0.029                      | 0.013                      | 0.013                                | 0.016                     |
| Number of id   |                             |                                       | 62                         |                            |                                      | 34                        |

**Note:** The dependent variable in columns (1)–(3) is the estimate for “*Bothin*”. Column (1) includes all estimates of this type. Column (2) only includes those that were published in a journal. Column (3) includes fixed effects by study. The value reported for the constant is the average of all fixed effects. The dependent variable in columns (4)–(6) is the estimate for “*Onein*”. Column (4) includes all estimates of this type. Column (5) only those that were published in a journal. Column (6) includes fixed effects by study. The value reported for the constant is the average of all fixed effects. Standard errors (in parentheses) are clustered at the study level and bootstrapped with 10,000 replications.

## 5 The impact of estimation characteristics

There are several potential candidates for explaining the heterogeneity in the estimates of the “WTO trade effect”, related to the methodology and data used. Gravity theory has developed rapidly over the last twenty years, so many studies may have used different specifications to estimate the gravity model. In addition, studies may focus on different samples, and there are many reasons to explain why there might be different WTO trade effects.

In light of recent advances in gravity theory, the literature suggests that certain features should be included when estimating a structural gravity equation. First, in line with Anderson and van Wincoop (2003), gravity models should account for multilateral resistance terms (MRTs), i.e., the fact that trade between two countries depends on the trade costs that these two countries face when trading with the rest of the world. Head and Mayer (2014), Yotov et al. (2016), and Larch et al. (2025b) suggest that a theory-consistent way to do this empirically is to include exporter-time and importer-time fixed effects. However, many studies do not account for MRTs or do so using atheoretical approaches.

Second, as suggested by Santos Silva and Tenreyro (2006), it is advisable to estimate the gravity equation in its multiplicative form using the Poisson pseudo-maximum likelihood (PPML) estimator. The PPML estimator has the advantage of accounting for heteroscedasticity and zero trade flows. However, various

regressions use the log-linearized transformation of the multiplicative trade equation implied by modern trade theory, and in some cases do not account for the loss of information by excluding observations with zeros, which are automatically excluded from the regressions by calculating the logarithm of the dependent variable.

Third, Baier and Bergstrand (2007) argue that trade policy variables in gravity models may be endogenous. For example, countries that trade more with each other may be more likely to sign a trade agreement. To mitigate such bias, they suggest including pair fixed effects, a practice that has become standard in the literature. Many studies use standard gravity variables instead, such as those that measure bilateral distance, identify a common language, or identify a common border.

Fourth, Yotov (2022) summarizes many theoretical and empirical reasons why the inclusion of domestic trade flows in gravity models is advisable, primarily because it aligns the empirical specification with gravity theory. Although the rationale for using domestic trade has spread rapidly in the literature, it is not yet a standard practice shared by all researchers or practitioners. The sample of countries used is also important. Advanced economies, on the one hand, and emerging and developing economies, on the other, may benefit differently from WTO membership, given their different economic and trade structures and capacities.

To study the effect of these different choices in estimation, we code six variables that describe the characteristics of the methodology and data used, and include these variables as additional explanatory variables in Equation (3). These indicators consist of a dummy variable indicating the absence of the use of multilateral resistance terms (MRTs) or the use of atheoretical ones,<sup>11</sup> a dummy variable for an estimation method other than PPML, a dummy variable for the absence of pair fixed effects, a dummy variable for the absence of domestic trade, and dummy variables for when the regression is on a sample in which both are AEs, or one is an AE and the other is not (most coefficients come from regressions that have no restriction of this form).

Tables 3 and 4 summarize the impact of different ways of estimating gravity equations on the WTO effect. Table 3 shows the results for the coefficient of *Bothin*, and Table 4 shows the results for the coefficient of *Onein*. The overarching conclusion from this analysis is that some of the estimation characteristics can have much larger effects than the potential publication bias in this case.

Among the variables considered, the most important negative effects on the coefficients for *Onein* and

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<sup>11</sup>They are usually approximated by the so-called “remoteness indexes”, computed as GDP-weighted distance averages.

Table 3: Estimation characteristics (*Bothin*)

| VARIABLES               | (1)      | (2)      | (3)      | (4)      | (5)       | (6)      | (7)       |
|-------------------------|----------|----------|----------|----------|-----------|----------|-----------|
| Standard error          | -0.314*  | -0.295   | -0.316*  | -0.309   | -0.313*   | -0.310*  | -0.291    |
|                         | (0.189)  | (0.189)  | (0.190)  | (0.193)  | (0.188)   | (0.184)  | (0.214)   |
| No or atheoretical MTRs |          | 0.080    |          |          |           |          | 0.058     |
|                         |          | (0.077)  |          |          |           |          | (0.073)   |
| Other than PPML         |          |          | 0.237*** |          |           |          | 0.240***  |
|                         |          |          | (0.066)  |          |           |          | (0.069)   |
| No pair FE              |          |          |          | -0.056   |           |          | -0.089*   |
|                         |          |          |          | (0.056)  |           |          | (0.046)   |
| No domestic trade       |          |          |          |          | -0.166*** |          | -0.351*** |
|                         |          |          |          |          | (0.051)   |          | (0.072)   |
| One advanced            |          |          |          |          |           | -0.091   | 0.001     |
|                         |          |          |          |          |           | (0.113)  | (0.082)   |
| Both advanced           |          |          |          |          |           | 0.065    | 0.057     |
|                         |          |          |          |          |           | (0.106)  | (0.089)   |
| Constant                | 0.258*** | 0.202*** | 0.082    | 0.289*** | 0.422***  | 0.261*** | 0.429***  |
|                         | (0.041)  | (0.068)  | (0.062)  | (0.050)  | (0.040)   | (0.037)  | (0.054)   |
| Observations            | 1,422    | 1,422    | 1,422    | 1,422    | 1,422     | 1,422    | 1,422     |
| R-squared               | 0.019    | 0.024    | 0.059    | 0.022    | 0.021     | 0.023    | 0.073     |

**Note:** The dependent variable is the estimate for “*Bothin*”. Standard errors (in parentheses) are clustered at the study level and bootstrapped with 10,000 replications.

*Bothin* come from not using domestic trade, meaning that the omission of domestic trade leads to lower coefficients on average in the sample of existing estimates. This finding is consistent with Dai et al. (2014), whose results indicate that estimates of trade policy variables are biased downward in regressions based on a sample that is limited to international trade only. In the case of *Onein*, the coefficients related to a sample consisting exclusively of AEs are also significantly larger than those from regressions without this restriction. The last column shows that the estimate of the WTO effects is significantly larger using the current best practices in estimation. The constant in the last column should be interpreted as the average effect coming from regressions in which theoretical MRTs are used, zero trade flows are not excluded, the PPML estimator is used, pair fixed effects are used to control for endogeneity, domestic trade data are available, and the sample does not select only AEs. Such a regression yields point estimates for the two coefficients of interest that are positive and substantially larger than the average estimates in the literature.<sup>12</sup>

<sup>12</sup>In further robustness tests, reported in Tables A.8 and A.9 in the Appendix, we run the same regressions on a sample winsorized at the 1% level to ensure that results are not driven by outliers. Results are in line with those reported in the main text. The relatively larger coefficient on *Onein* in Columns 5 and 7 should be interpreted with caution. In our sample, these estimates are driven by a single study available only in its working paper version, which limits the robustness of the inference.

Table 4: Estimation characteristics (*Onein*)

| VARIABLES               | (1)                | (2)               | (3)                | (4)                 | (5)                  | (6)                | (7)                  |
|-------------------------|--------------------|-------------------|--------------------|---------------------|----------------------|--------------------|----------------------|
| Standard error          | 0.002<br>(0.312)   | 0.002<br>(0.400)  | 0.002<br>(0.311)   | 0.002<br>(0.309)    | 0.002<br>(0.317)     | 0.002<br>(0.319)   | 0.002<br>(0.357)     |
| No or atheoretical MTRs |                    | -0.106<br>(0.228) |                    |                     |                      |                    | 0.046<br>(0.282)     |
| Other than PPML         |                    |                   | -0.207<br>(0.151)  |                     |                      |                    | -0.036<br>(0.107)    |
| No pair FE              |                    |                   |                    | -0.160**<br>(0.071) |                      |                    | -0.207***<br>(0.063) |
| No domestic trade       |                    |                   |                    |                     | -0.486***<br>(0.048) |                    | -0.369<br>(0.307)    |
| One advanced            |                    |                   |                    |                     |                      | 0.355**<br>(0.173) | 0.434*<br>(0.240)    |
| Both advanced           |                    |                   |                    |                     |                      | 0.293**<br>(0.127) | 0.318**<br>(0.139)   |
| Constant                | 0.138**<br>(0.058) | 0.235<br>(0.233)  | 0.338**<br>(0.146) | 0.245***<br>(0.056) | 0.618***<br>(0.051)  | 0.088<br>(0.058)   | 0.575***<br>(0.065)  |
| Observations            | 862                | 862               | 862                | 862                 | 862                  | 862                | 862                  |
| R-squared               | 0.013              | 0.016             | 0.018              | 0.033               | 0.023                | 0.064              | 0.108                |

**Note:** The dependent variable is the estimate for “*Onein*”. Standard errors (in parentheses) are clustered at the study level and bootstrapped with 10,000 replications.

## 6 Gravity estimates of the WTO trade effect

We now turn to estimating the impact of the WTO on bilateral trade using a theory-consistent version of the gravity specification. In this section, we focus specifically on examining the impact of the WTO among its member countries (i.e., the *Bothin* coefficient in the previous section) and relegate the effect between members and non-members (*Onein*) to the end of the section. We use a standard empirical specification (e.g., Yotov et al., 2016), which takes the form:

$$X_{ijt} = \exp(\beta_1 WTO_{ijt} + \psi Z'_{ijt} + \eta_{it} + \theta_{jt} + \omega_{ij}) + \varepsilon_{ijt}, \quad (4)$$

where the dependent variable  $X_{ijt}$  refers to gross bilateral trade flows between the exporter  $i$  and importer  $j$  in year  $t$ , also including the case where  $i = j$  (i.e., domestic trade flows). As discussed in the literature (Dai et al., 2014; Yotov, 2022; Larch et al., 2025a), the inclusion of both international and domestic trade flows is consistent with trade theory.<sup>13</sup> The variable  $WTO_{ijt}$  is a dummy variable, equal

<sup>13</sup>In fact, we are not aware of any theoretical international trade model that would exclude domestic sales from the analysis.

to 1 when countries  $i$  and  $j$  are both members of the WTO at time  $t$ , and zero otherwise. Thus, the coefficient  $\beta_1$  captures the semi-elasticity of bilateral trade flows for WTO membership. Importantly, our strategy for identifying the “WTO effect” closely mirrors the approach of the literature analyzed in the meta-analysis. This means that the dummy variable  $WTO_{ijt}$  captures both changes in tariffs and non-tariff measures associated with WTO membership.

The vector  $Z'_{ijt}$  contains control variables that capture other time-varying trade policies. In our main specification, this vector includes dummy variables identifying membership in trade agreements ( $TA_{ijt}$ ) and the European Union ( $EU_{ijt}$ ). As a robustness check, we further modify and expand the vector  $Z'_{ijt}$ , for example, by allowing the effect of bilateral distance to vary over time, an econometric strategy often interpreted as a proxy for bilateral transport costs. The terms  $\eta_{it}$  and  $\theta_{jt}$  are exporter-time and importer-time fixed effects, respectively. They are important for two main reasons: first, they represent the theory-consistent way to control for “multilateral resistance terms” (Anderson and van Wincoop, 2003); second, they control for all features having country-time variation (such as GDP, GDP per capita, population, etc.). The term  $\omega_{ij}$  represents directional (i.e., depending on the direction of trade) country-pair fixed effects, controlling for all features with exporter-importer (pair) variation (such as the standard gravity variables: distance, contiguity, common language, colonial relationship, etc.). It is also a standard way of accounting for all (observable and unobservable) time-invariant bilateral components of trade costs.  $\varepsilon_{ijt}$  is the error term.

When gravity is estimated with domestic trade flows, it is a standard practice to include a series of time-varying dummy variables that distinguish international trade flows from domestic trade flows. The coefficients of these variables are usually interpreted as proxies for trade globalization because they capture the ease of trading internationally relative to trading domestically (Bergstrand et al., 2015). This avoids attributing general changes in the ease of trading across international borders brought about by multilateral progress to other bilateral or regional trade policy variables, such as trade agreements. However, as our primary focus is on estimating the WTO effect, this can be problematic.

Unlike the so-called first globalization (1870–1913), where the ease of trading internationally was mainly driven by falling transport costs (see, e.g., Jacks et al., 2010), globalization in the post-World War II period (our period of analysis) was mainly driven by falling trade policy-related costs (see, e.g., Anderson and van Wincoop, 2004), such as multilateral tariff reductions promoted by the WTO. In this case, including the trade globalization proxy will remove some of the WTO effects from the WTO variable.

There is no simple solution to this problem. Therefore, we interpret the estimates from specifications that do not explicitly use time-varying border variables to control for common globalization effects as an upper-bound estimate of the WTO effect. In addition, however, we also estimate a version of this equation with the addition of a series of time-varying dummy variables as follows:

$$X_{ijt} = \exp(\beta_1 WTO_{ijt} + \phi_t b_{ij} + \psi Z'_{ijt} + \eta_{it} + \theta_{jt} + \omega_{ij}) + \varepsilon_{ijt}. \quad (5)$$

The only difference to the previous equation is the appearance of  $\phi_t b_{ij}$ , where the variable  $b_{ij}$  is a dummy variable that distinguishes international trade flows ( $b_{ij} = 1$ ) from domestic trade flows ( $b_{ii} = 0$ ), and the time-varying coefficients  $\{\phi_t\}$  capture the semi-elasticity of bilateral trade flows to crossing an international border in different years. Following our previous discussion, we interpret the specification of equation (5) as a lower-bound estimate of the WTO effect.

We estimate specifications (4) and (5) using a Poisson Pseudo-Maximum Likelihood (PPML) estimator, which, differently from OLS, addresses two key characteristics of the trade data: (i) heteroskedasticity, and (ii) the zero trade flows (Santos Silva and Tenreyro, 2006). We apply this estimation procedure to seven different databases. This multi-database approach ensures that our conclusions are not driven by the choice of a single dataset and enables us to leverage the unique strengths of each database, whether broader country and sector coverage, longer time spans, or alternative measures of domestic trade. Some databases allow us to capture early entries into the WTO system, analyze accession patterns across both AEs and EMDEs, and retain a large pool of untreated bilateral pairs for comparison. Others provide theory-consistent measures of domestic trade flows, calculated as the difference between nominal gross production and nominal total exports.<sup>14</sup> In this sense, our approach is close to the spirit of meta-analysis:

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<sup>14</sup>We calculate total exports as the sum of bilateral exports. In several databases, domestic trade can only be constructed as nominal GDP minus nominal total exports. Despite theoretical concerns about value added versus gross measures, Campos et al. (2021) compare gravity models using GDP-based domestic trade with those using gross output and find that estimated coefficients for trade policy variables like trade agreements are very similar across approaches. They argue that theory-consistent fixed effects (exporter-time, importer-time, and pair fixed effects) in structural gravity equations absorb systematic differences between GDP and gross output, limiting the practical relevance of the domestic trade measure. When domestic trade is proxied using GDP-based measures, a small share of flows appears negative. We treat these as data inconsistencies and replace them with missing values. These negative flows account for approximately 1% to 2% (depending on the database) of domestic trade observations. Gross output-based databases include only manufacturing output and trade, whereas GDP-based databases include all merchandise trade. For databases that approximate domestic trade using GDP-based measures, we conduct a series of robustness tests (reported in Appendix B) employing an alternative, output-based, database for domestic trade developed in Head and Mayer (2021). The database is publicly available and covers the period 1950-2018. The authors reconstruct country-year production values for four goods sectors (agriculture, fishing, mining, and manufacturing) and then subtract total goods exports. Since aggregate production of goods is not consistently available, the authors combine multiple sources and imputation methods. They use OECD, UNIDO, and UN national accounts data to obtain industry value added and output, and where output is missing, they apply stable output-to-value added ratios derived from OECD benchmarks. Manufacturing is treated with a hierarchy of sources: UNIDO output data when available, OECD data as a second option, and interpolated output-to-GDP ratios if the

rather than privileging a single dataset or measure, we apply a common specification across different sources and assess the consistency of results across the full range of databases used.

The seven databases that we employ are the following: (i) merchandise trade flows from the International Monetary Fund’s (IMF) Direction of Trade Statistics (DOTS) as reported by the country of destination (1948-2019); (ii) merchandise trade flows from IMF DOTS as reported by the country of origin (1948-2019); (iii) merchandise trade flows from the United Nations Comtrade database as reported by the country of destination (1962-2019); (iv) merchandise trade flows from United Nations Comtrade database as reported by the country of origin (1962-2019); (v) merchandise trade flows from CEPII BACI (1996-2019, Gaulier and Zignago, 2010);<sup>15</sup> and (vi) manufacturing trade flows from CEPII TradeProd (1966-2018, de Sousa et al., 2012; Mayer et al., 2023); (vii) manufacturing trade flows from the WTO Structural Gravity Database (1980-2016; Larch et al., 2025a). Trade flows are gross, nominal, and expressed in a common currency (thousands of US dollars). In the main text, we report the median estimate across databases, however, we report the results from all databases in the Appendix. Finally, the data on WTO membership also come from the CEPII gravity database (Head et al., 2010).

**Deviations from standard practice.** In Figure 4, we show the point estimates and confidence intervals for the two alternative specifications (with and without border-time dummies) obtained from estimating a theory-consistent structural gravity model.<sup>16</sup> Appendix B. Reassuringly, across all theory-consistent specifications, both the TA and EU coefficients are positive. Moreover, the EU coefficient consistently exceeds the WTO and TA coefficients. This pattern reflects the depth of trade integration within the EU, which is substantially greater than that of any other integration initiative worldwide. This observation aligns with findings in recent literature, such as Head and Mayer (2021) and Nagengast et al. (2026). We also show the results obtained when one deviates from standard practice, such as using OLS regression on log trade, not controlling for MRTs, not including pair fixed effects, or not considering domestic trade flows. Both the “upper bound” and “lower bound” equations yield similar patterns across specifications.

The median upper-bound estimate yields a point estimate of 0.41, implying an increase in trade flows of  $\exp(0.41) - 1 = 51\%$ . The median lower-bound estimate yields a point estimate of 0.30 and an increase

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previous two options are not available. Summing the reconstructed outputs across the four tradable sectors yields total goods production, from which exports—sourced mainly from UN Comtrade and IMF DOTS—are subtracted to obtain domestic trade. This procedure produces a coherent measure of domestic tradable goods absorption that can be compared systematically across countries and time.

<sup>15</sup>We access IMF DOTS, United Nations Comtrade, and BACI trade flows via the Centre d’Etudes Prospectives et d’Informations Internationales (CEPII) gravity database (Head et al., 2010).

<sup>16</sup>We report the point estimates and confidence intervals for the WTO, trade agreements (TA), and the European Union (EU) variables included in the regression in

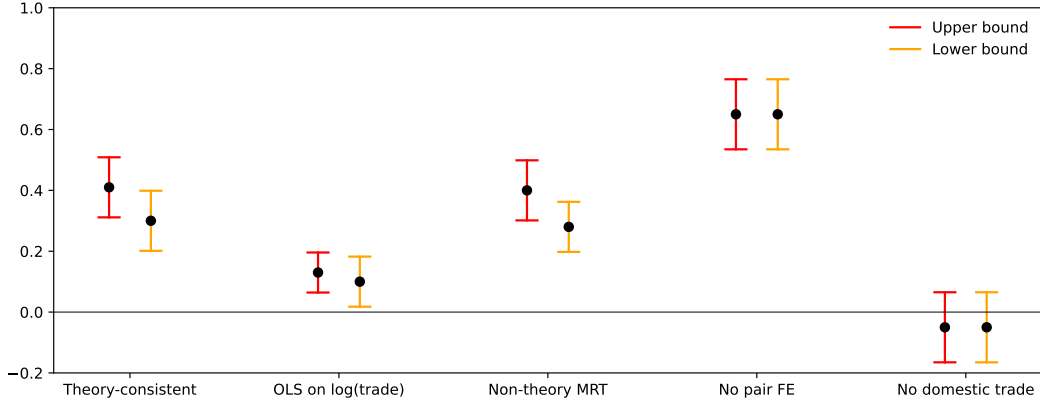


Figure 4: Deviations from standard practice

**Note:** The figures show point estimates (black dots) and 90% confidence intervals for the “upper bound” (in red) and for the “lower bound” (in yellow). Standard errors are clustered by country pair. We report the median value for estimations based on seven different databases. The x-axis of the figure lists the alternative specifications, and highlights the characteristic that deviates from standard practice (or “theory-consistent” specification, whose upper and lower bound coefficients are estimated using Equation 4 and Equation 5, respectively). For example, the “OLS on log(trade)” means that we are using an OLS estimator instead of PPML, while keeping the other features equal to the “theory-consistent” specification; “no MRT” means that we are not using exporter-time and importer-time fixed effects, while keeping the other features equal to the “theory-consistent” specification; etc..

in trade flows of  $\exp(0.30) - 1 = 35\%$ . We note a substantial effect of not including pair fixed effects on the WTO coefficient. This finding is consistent with Baier and Bergstrand (2007), highlighting the importance of using a panel approach with country-pair fixed effects to fully control for all time-invariant bilateral trade costs and mitigate endogeneity concerns, which otherwise is likely to lead to an upward bias in the coefficient estimates. We also note the key role of domestic trade in the empirical estimates, the exclusion of which can lead to a downward bias in the estimated coefficient of the trade policy variables, as demonstrated for the WTO effects by Larch et al. (2025a).

**Robustness tests.** We test the robustness of our preferred specification by allowing the effect of bilateral distance to vary over time or by testing alternative clustering strategies, as suggested by Egger and Tarlea (2015). The former allows capturing the effects of policy and non-policy variables, such as the construction of road or rail infrastructure connecting (or improving) two countries. The results are consistent with our main specification and are shown in Figure 5.<sup>17</sup>

**Country income levels.** We next examine whether the large, positive, and significant effect of WTO membership is heterogeneous across country income levels. To do this, we interact our main variable of

<sup>17</sup>Unsurprisingly, allowing the effect of bilateral distance to vary over time has only a marginal impact when the specification already accounts for time-varying international borders.

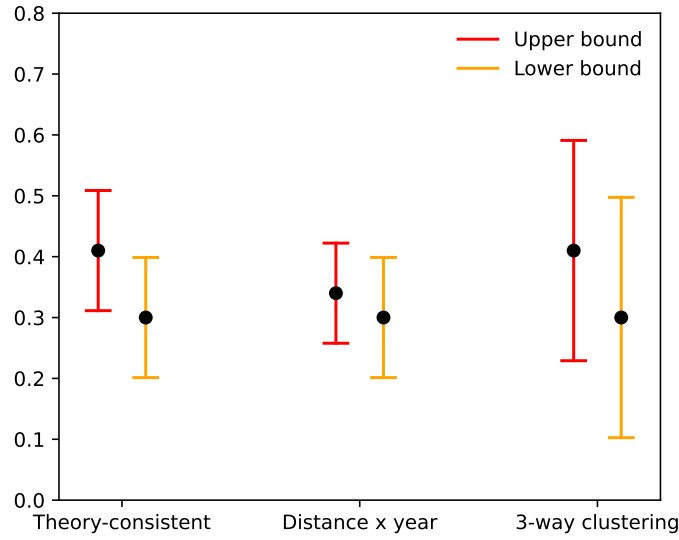


Figure 5: Robustness

**Note:** The figures show point estimates (black dots) and 90% confidence intervals for the “upper bound” (in red) and for the “lower bound” (in yellow). ‘Distance x year’ refers to a specification that adds interactions of the logarithm of distance and year dummies to the covariates. Unless otherwise indicated, standard errors are clustered by country pair. ‘3-way clustering’ refers to clustering by exporter, importer, and time. We report the median value for estimations based on seven different databases.

interest ( $WTO_{ijt}$ ) with four different dummy variables that capture whether the trade flow is between two advanced economies (AEs-AEs), between an advanced economy and an emerging economy (also taking into account who is the exporter and who is the importer; AEs-EMDEs and EMDEs-AEs), or between emerging markets and developing economies (EMDEs-EMDEs).

The estimates in Figure 6 show that the WTO effect is largely concentrated in AE-EMDE and EMDE-EMDE trade flows. These results support the idea that the reductions in trade barriers, such as tariffs and non-tariff barriers, achieved and implemented through successive rounds of trade negotiations have promoted the integration of EMDEs into global value chains, while at the same time allowing AEs to gain access to lower-cost inputs. This mutually beneficial relationship has enabled both AEs and EMDEs to expand their exports to each other. At the same time, many EMDEs (such as China, India, or Brazil, to name a few) have joined or deepened their participation in the global trading system in recent decades. By doing so, they also promote trade among EMDEs by strengthening regional value chains based on WTO integration.

On the other hand, trade among AEs does not seem to have been significantly affected by WTO membership. This may reflect two complementary factors. First, other trade integration initiatives (such

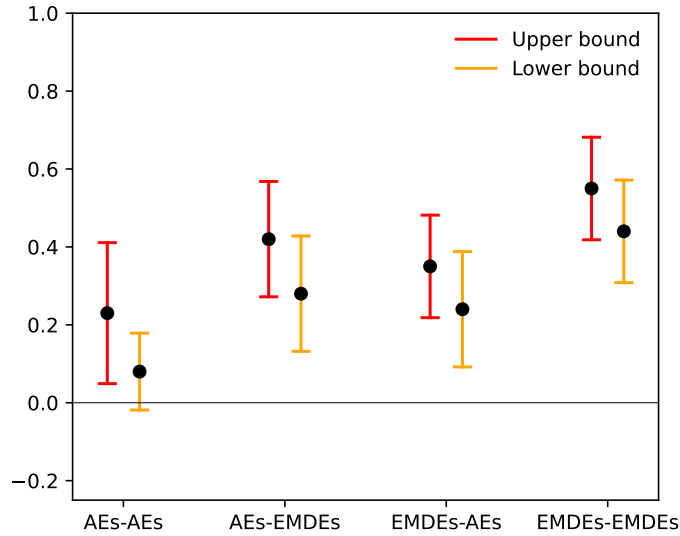


Figure 6: Results by country income level

**Note:** The figures show point estimates (black dots) and 90% confidence intervals for the “upper bound” (in red) and for the “lower bound” (in yellow). Standard errors are clustered by country pair. We report the median value for estimations based on seven different databases. The label ‘AEs’ refers to ‘Advanced Economies’. The label ‘EMDEs’ refers to ‘Emerging Market and Developing Economies’, as classified by the IMF.

as the European Union), which have further promoted integration among AEs, may have played a more important role in this case. Second, the WTO trade effects for trade among AEs likely occurred during the early GATT era, when most of the AEs joined the multilateral system and liberalized their markets, but not, on average, over a long sample.<sup>18</sup>

**Sectors.** We also examine whether the WTO effect is heterogeneous across sectors. For this purpose, we use the OECD TiVA database, which contains consistent international and domestic trade data for 76 economies from 1995 to 2019. The OECD TiVA database provides bilateral trade data for 45 industries that can be consistently aggregated, as we do, into four main economic sectors: agriculture, mining, manufacturing, and services.<sup>19</sup> Figure 7 shows estimates of the WTO effect on bilateral trade flows between members for the four main economic sectors. The main message from this figure is that the

<sup>18</sup>To test this hypothesis, we divided the sample into three periods of similar length: an “early GATT” period (1948–1973), a “late GATT” period (1974–1994), and a “WTO period” (1995–onward). The results, reported in Table B.4 in the Appendix, indicate that the WTO effect on trade among advanced economies was concentrated in the early GATT period and subsequently extended to other trade flow combinations.

<sup>19</sup>For a more granular analysis, we exploit the ITPD-E database (The International Trade and Production Database for Estimation), which contains consistent data on international and domestic trade for industries within the agriculture, mining, energy, manufacturing, and services sectors (1986–2019 for agriculture; 1988–2019 for mining, energy, and manufacturing; 2000–2019 for services; see Borchert et al. (2021) for more details). Results, reported in Figure B.1 using this database also show that the WTO trade effects are large and positive, but very heterogeneous across sectors and industries within sectors.

WTO has had a large, positive, and significant effect on trade in goods and trade in services, although the effect tends to be larger for the former group. The effect of the WTO on agricultural and manufacturing trade flows is about two to three times larger than its effect on services.

**Country income levels and sectors.** Figure 8 shows sector estimates that allow for heterogeneity in the income levels of trading partners. According to our main estimates, WTO membership favours more exports from EMDEs in agriculture (both to AEs and other EMDEs). In mining and energy, WTO membership favours more exports from AEs to EMDEs. In manufacturing, WTO membership favours exports across the board, i.e., across all levels of development (i.e., AEs-AEs, AEs-EMDEs, EMDEs-AEs, and EMDEs-EMDEs). For services, the results are qualitatively similar to those for manufacturing, but the effects are smaller.<sup>20</sup>

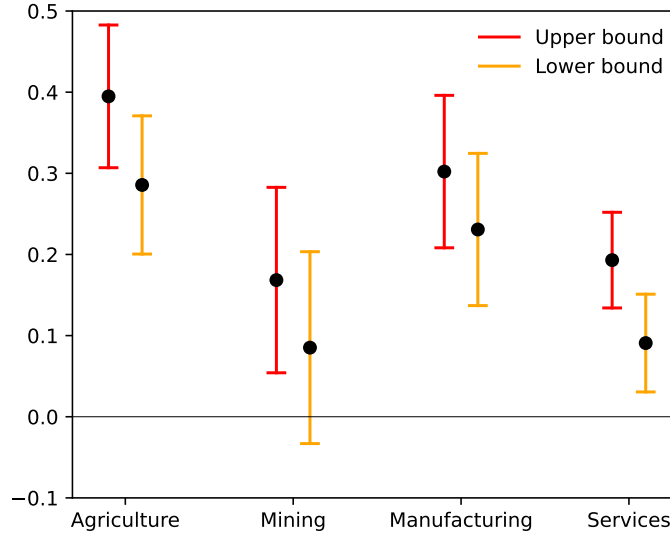


Figure 7: Results by sector

**Note:** The figures show point estimates (black dots) and 90% confidence intervals for the “upper bound” (in red) and for the “lower bound” (in yellow). Standard errors are clustered by country pair. The sector labelled mining also includes energy. Estimations use data from the OECD TiVA database.

**Separating GATT from WTO.** We also distinguish between the effects of GATT and WTO. The coefficients obtained for GATT and WTO separately, as shown in Table B.5 in the Appendix, are similar in magnitude across all specifications. However, the coefficient for the WTO is estimated with greater precision, possibly due to the larger number of trade flows and entries documented in our databases for

<sup>20</sup>As the analysis draws on the OECD TiVA database, which covers the period from 1995 to 2019, the set of AEs used to estimate the AEs-AEs effect is necessarily limited. Only a few AEs (Estonia, Latvia, and Lithuania) joined the WTO during this time frame, limiting the variation available for identifying this effect in this specific exercise.

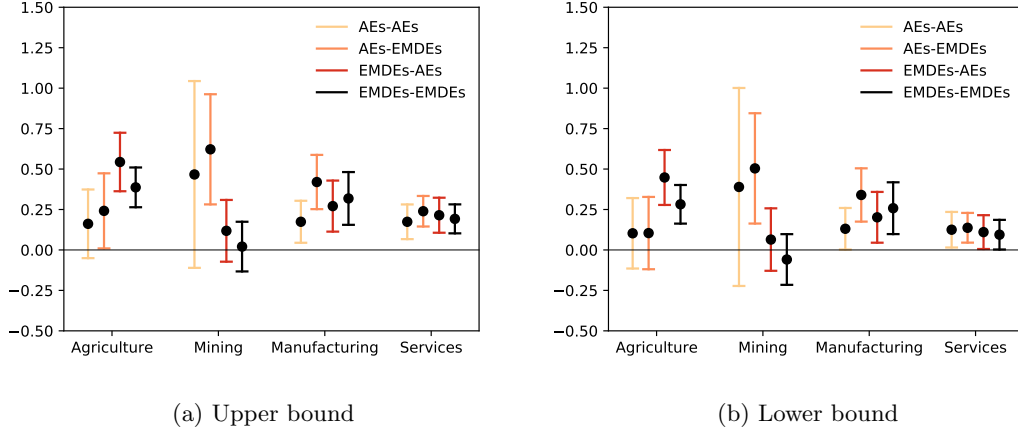


Figure 8: Results by sector and country income level

**Note:** The figures show point estimates (black dots) and 90% confidence intervals (vertical error bars). Standard errors are clustered by country pair. The sector labelled mining also includes energy. Estimations use data from the OECD TiVA database. The label ‘AEs’ refers to ‘Advanced Economies’. The label ‘EMDEs’ refers to ‘Emerging Market and Developing Economies’, as classified by the IMF. The four estimations on the left do not control for border-year dummies (“upper bound”). The four estimations on the right control for border-year dummies (“lower bound”).

the period during which the WTO was active.

**Bothin and Onein.** We then extend our analysis to assess the impact of the WTO on trade flows between members and non-members. WTO members sometimes reduce trade barriers against all countries, not just other WTO members. In our model, as shown by Heid et al. (2021), it is possible to identify the variable *Onein* even when using the full set of exporter-time and importer-time fixed effects, because domestic trade flows are included.<sup>21</sup> Our estimates (reported in Table 5) show that the WTO trade effect tends to be positive for trade flows between members and non-members. Moreover, our results consistently indicate that the magnitude of the WTO trade effect for trade flows between members and non-members is smaller than the effect observed for trade flows between members. This suggests that non-members do not take full advantage of the benefits of WTO membership and that only part of the WTO-related reductions in trade barriers are granted to all. If anything, our findings support the notion that the WTO is primarily a trade-creating institution rather than a trade-diverting one. These results are consistent with previous studies on this topic, which have found a positive impact of the WTO on

<sup>21</sup>The variables *Bothin* and *Onein* are not collinear with exporter-time and importer-time fixed effects because they equal zero for domestic trade flows. This is equivalent to interacting the variable *Onein* with the variable  $b_{ij}$ , used in Equation 5, a dummy variable that distinguishes international trade flows ( $b_{ij} = 1$ ) from domestic trade flows ( $b_{ii} = 0$ ). The variable thus specified captures the differential effect of *Onein* on international trade compared to domestic trade. However, as shown in Campos et al. (2023) and Beverelli et al. (2024), this strategy does not allow to separately identify the cases where the exporter is a WTO member from those where the importer is, while maintaining all fixed effects.

trade flows between members and non-members (Subramanian and Wei, 2007; Larch et al., 2025a).

To this extent, our results suggest two important implications for the measurement of WTO effects in gravity models when *Bothin* and *Onein* are both included in the regression. First, when domestic trade flows are available, and the *Onein* variable is not included, the *Bothin* effect is likely to be biased downwards. This bias arises because *Bothin* and *Onein* are negatively correlated by construction and *Onein* seems to increase when a country joins the WTO. Or, in other words, when not including *Onein*, the reference group for *Bothin* contains all trade flows where neither country is a WTO member and all trade flows where one country is a WTO member alongside domestic sales. When the *Onein* effect is positive, this suggests that, on average, trade flows of the reference group are also increasing, leading to a smaller estimate for the *Bothin* coefficient. Second, when both *Bothin* and *Onein* are included in the regression, the identification of both effects also relies strongly on the evolution of domestic trade flows. In this case, estimates are sensitive to whether domestic trade has been constructed using GDP-based proxies or gross output data.<sup>22</sup> Given this concern, and the use of many different databases, we adopt an “incremental approach”, first estimating a specification that include only *Bothin*, and then extending the analysis to jointly include *Bothin* and *Onein*. This allows us to explain and decompose any differences in the estimated coefficients. We acknowledge that this choice is conservative, but it provides a more reliable benchmark by avoiding the sensitivity associated with jointly identifying *Bothin* and *Onein*.

Table 5: *Bothin* versus *Onein*

|                        | Upper bound |              | Lower bound |              |
|------------------------|-------------|--------------|-------------|--------------|
|                        | Both in     | One in       | Both in     | One in       |
| IMF DOTS (importer)    | 0.71 (0.13) | 0.28 (0.11)  | 0.56 (0.13) | 0.23 (0.11)  |
| IMF DOTS (exporter)    | 0.81 (0.13) | 0.35 (0.12)  | 0.62 (0.13) | 0.29 (0.11)  |
| UN COMTRADE (importer) | 0.68 (0.15) | 0.28 (0.13)  | 0.57 (0.14) | 0.26 (0.12)  |
| UN COMTRADE (exporter) | 0.87 (0.17) | 0.49 (0.16)  | 0.75 (0.16) | 0.46 (0.15)  |
| BACI CEPII             | 0.35 (0.12) | 0.10 (0.10)  | 0.28 (0.12) | 0.08 (0.09)  |
| TradeProd CEPII        | 0.29 (0.12) | -0.01 (0.09) | 0.11 (0.12) | -0.04 (0.09) |
| WTO Structural DB      | 0.59 (0.20) | 0.05 (0.19)  | 0.20 (0.20) | -0.05 (0.19) |

**Note:** The table reports the point estimates and standard errors for the variables *Bothin* and *Onein*. Standard errors are clustered by country pair. Results in the first two columns do not control for border-year dummies (“upper bound”). Results in the last two columns control for border-year dummies (“lower bound”). The rows correspond to various databases used in the estimation.

<sup>22</sup>This is consistent with Campos et al. (2021). Their primary finding is that, when comparing datasets using GDP-based domestic trade with those using gross output, the estimated coefficients for trade policy variables like trade agreements are very similar across approaches. However, a secondary finding is that, in cases where the identification of the policy variable relies solely on the variation between international and domestic trade, the estimates become sensitive to how domestic trade is constructed.

**Comparing meta-analysis findings and our gravity model estimates.** For *Bothin*, the average effect reported in the meta-analysis,  $\exp(0.21) - 1 = 23\%$  (Table 1), is lower than the median of our gravity regressions across different databases (between  $\exp(0.30) - 1 = 35\%$  and  $\exp(0.41) - 1 = 51\%$ , lower-bound and upper-bound specification, respectively), suggesting that the meta-analysis (partly based on older specifications) produces more conservative estimates. When focusing on meta-analysis results from specifications similar to ours,  $\exp(0.429) - 1 = 54\%$  (Column 7, Table 3), however, the size of the coefficient is broadly consistent with our findings.

For the *Onein* coefficient, the average effect from the meta-analysis,  $\exp(0.13) - 1 = 14\%$  (Table 1), is again lower than the median of our gravity regressions (between  $\exp(0.23) - 1 = 26\%$  and  $\exp(0.28) - 1 = 32\%$ , lower-bound and upper-bound specification, respectively), again suggesting that the meta-analysis (partly based on older specifications) produces more conservative estimates. However, in this case, the variation of our gravity estimates across databases is substantial, with some results even suggesting that the *Onein* effect is not statistically different from zero. This is driven by the fact that when both variables are included, identification relies heavily on the construction of domestic trade flows, which is sensitive to whether GDP-based proxies or gross output data are used. Finally, regarding the relative size of the coefficients, in nearly all cases, both in our regressions and in the meta-analysis, the *Bothin* effect is larger than the *Onein* effect. We find this result intuitive.<sup>23</sup>

**Contextualizing the size of the WTO coefficient.** To assess the economic magnitude of our estimated WTO coefficient, we translate it into an ad-valorem tariff-equivalent trade-cost wedge implied by WTO membership using a standard structural gravity framework. Under standard gravity, proportional changes in bilateral trade costs map into changes in bilateral trade flows, with the magnitude of this mapping governed by the trade elasticity parameter  $\theta$ .

Following standard practice in the literature, we adopt  $\theta = 5$  as a benchmark value for the trade elasticity parameter, consistent with the meta-analysis of Head and Mayer (2014). Using this elasticity, our preferred estimates of the WTO coefficient, based on databases covering the period from 1948 to 2019 (IMF DOTS), imply a reduction in bilateral trade costs equivalent to a tariff decline in the range of 6.6 to 9.9 percentage points.<sup>24</sup> Similar calculations, based on databases covering the period from 1996 to 2019 (BACI), imply a reduction in bilateral trade costs equivalent to a tariff decline in the range

<sup>23</sup>The sole exception arises in the meta-analysis of state-of-the-art regressions, where the *Onein* coefficient appears larger. Yet, this result is driven entirely by a single working paper, and should therefore be interpreted with caution.

<sup>24</sup>The ad-valorem equivalent (AVE) is computed as:  $AVE = 100 \times [\exp(\hat{\beta}/\theta) - 1]$ , where  $\hat{\beta}$  is the estimated coefficient and  $\theta$  denotes the trade elasticity.

of 4.1 to 5.1 percentage points. These tariff equivalent magnitudes can be compared to the observed evolution of average world tariff levels over time. According to Bown and Irwin (2015), average world tariffs stood at approximately 22% in 1947. Using data from Snoussi-Mimouni and Drevinskas (2023), average tariffs declined to 13.2% in 1996 and to 7.8% in 2019, the final year of our sample.<sup>25</sup>

Comparing these magnitudes, the tariff-equivalent effect associated with WTO membership accounts for up to 70% of the total decline in average world tariffs observed between 1947 and 2019, and for up to 94% of the tariff decline observed since the mid-1990s. Taken together, these calculations help anchor our regression estimates in historically observed changes in trade policy. Translating the WTO coefficient into tariff equivalent trade cost reductions yields magnitudes that represent a substantial share of the long-run decline in average world tariffs, particularly since the mid-1990s. A potentially important implication of this analysis is that the WTO's contribution to lowering trade costs has arisen predominantly from tariff reductions, with non-tariff barriers contributing to a lesser extent.

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<sup>25</sup>These figures refer to MFN applied tariffs, which most directly capture multilateral tariff commitments under the WTO. Using applied tariffs including preferences would confound WTO-related tariff changes with preferential liberalization outside the WTO framework. See Snoussi-Mimouni and Drevinskas (2023) for details on construction, coverage, and global trends in tariff data.

## 7 Conclusions

In this paper, we perform a quantitative review of the literature using meta-analysis techniques, complemented by direct empirical analysis of trade data based on theory-consistent structural gravity models, to assess the trade effect of the WTO. Specifically, we first collect 2,483 estimates of the WTO trade effect from 71 papers and construct variables characterizing the salient features of the methodology used for each estimate. We then assess the presence of publication bias and the impact of estimation characteristics on the estimated WTO trade effect. We also implement a theory-consistent gravity model to provide a theory-based update of the estimated WTO trade effects. In our preferred specification, we use both domestic and international trade flows, three-way fixed effects (exporter-time, importer-time, exporter-importer) to account for MRTs and trade policy endogeneity, and apply a Poisson pseudo-maximum likelihood estimator.

Our meta-analysis indicates that, on average, the WTO effect on trade between members reported in the literature is significantly positive, amounting to approximately +23%. We also find no evidence of publication bias, a finding that may be related to the existing debate in the literature. However, the reported estimates vary considerably and depend on the characteristics of the studies. In other words, how gravity equations are specified matters for trade cost estimates such as WTO membership. This conclusion should not be surprising given recent theoretical and empirical advances in the literature. Our structural gravity estimates confirm these findings: the WTO matters for trade. The median estimates from our lower- and upper-bound specifications run on seven different trade databases suggest that WTO membership is associated with a substantial and statistically significant increase in trade between members, ranging from +35% to +51%. These values are approximately 1.5 to 2 times the average effect reported in the meta-analysis. These results highlight the importance of well-specified gravity models in capturing the full WTO effect, and suggest that in this case, the meta-analysis, partly based on older specifications, tends to produce more conservative estimates. Drawing on two additional databases, we also find that the WTO effects are heterogeneous across sectors and income levels of trading partners.

Looking ahead, our evidence of a large but heterogeneous increase in trade among WTO members does not shed light on within-country effects associated with WTO membership (e.g., firm- or household-level effects). This issue merits further research and consideration.

## Data availability

Replication data for this article has been deposited at Campos et al. (2026)

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## A Meta-analysis

### A.1 Studies surveyed in the meta-analysis

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## A.2 Additional information

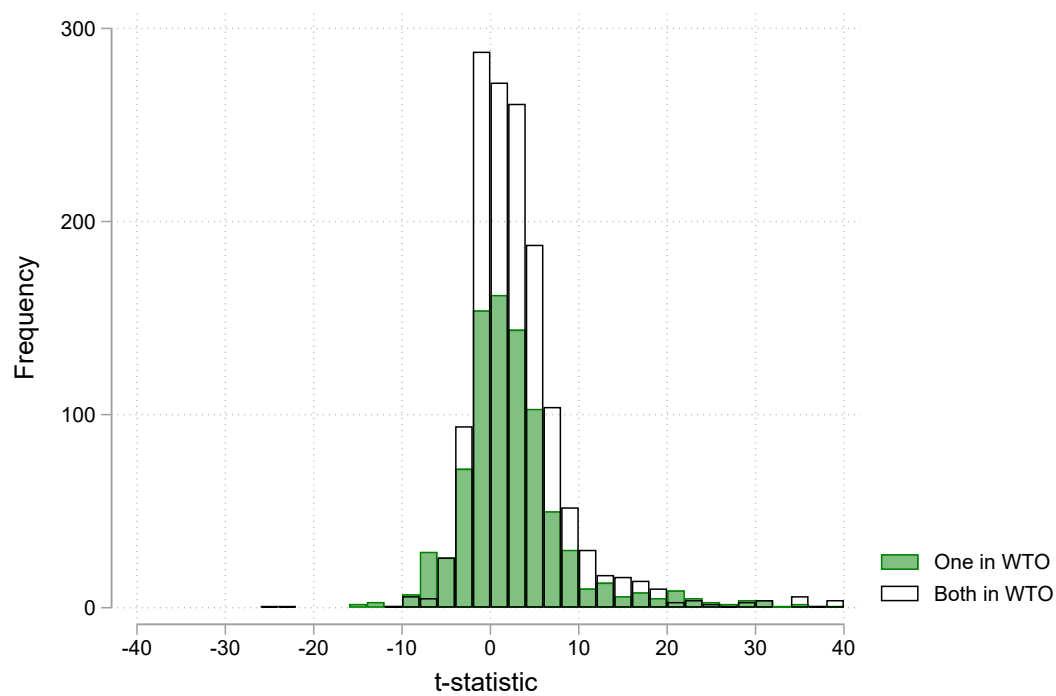


Figure A.1: Histogram of t-statistics

**Note:** The figure shows the histogram of the t-statistic of the WTO effect on trade in individual studies.

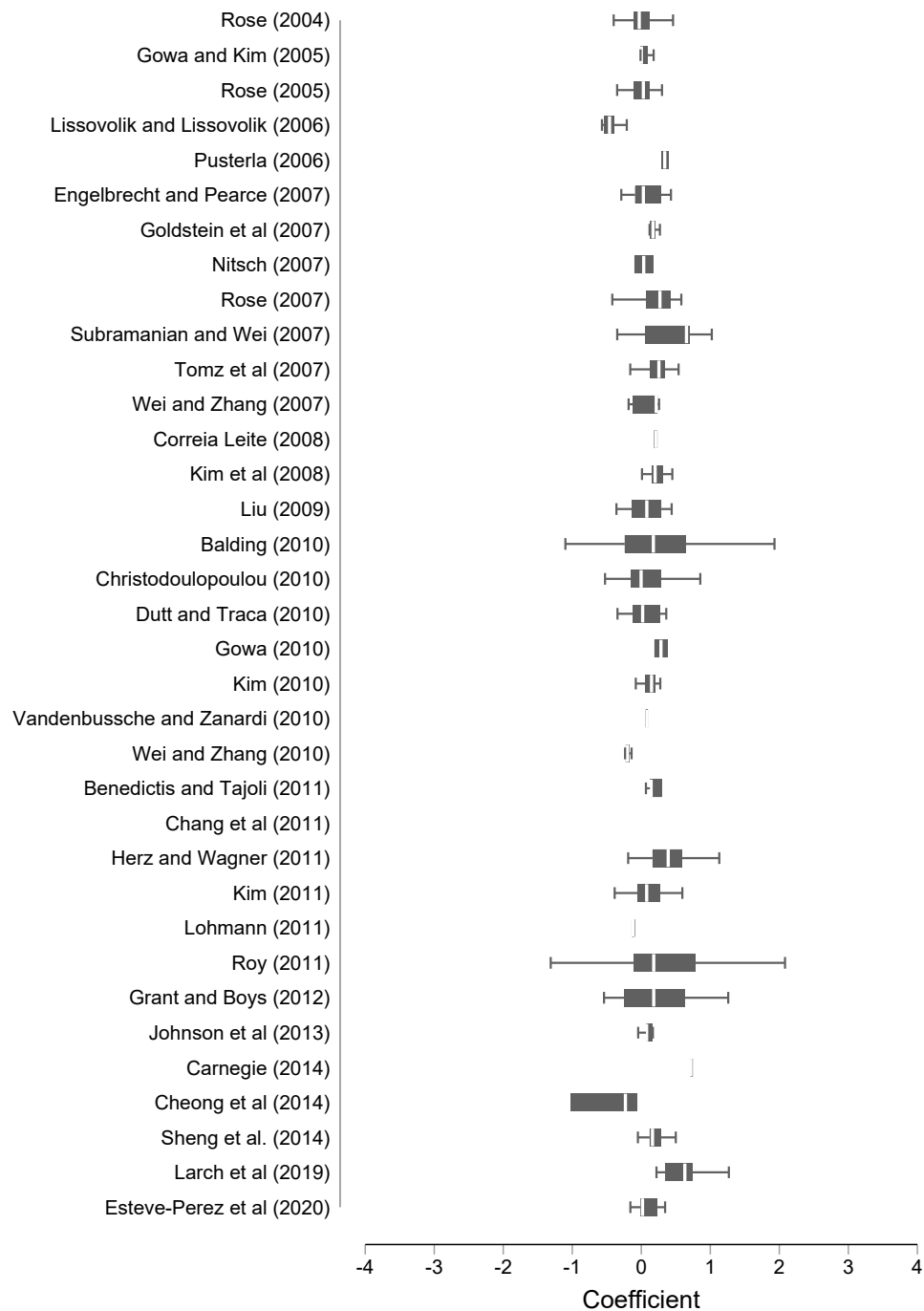


Figure A.2: Box plots for estimates for "One in WTO"

**Note:** Papers listed in chronological order. Outside values are not shown. Papers with too few observations to calculate percentiles are not removed from the list.

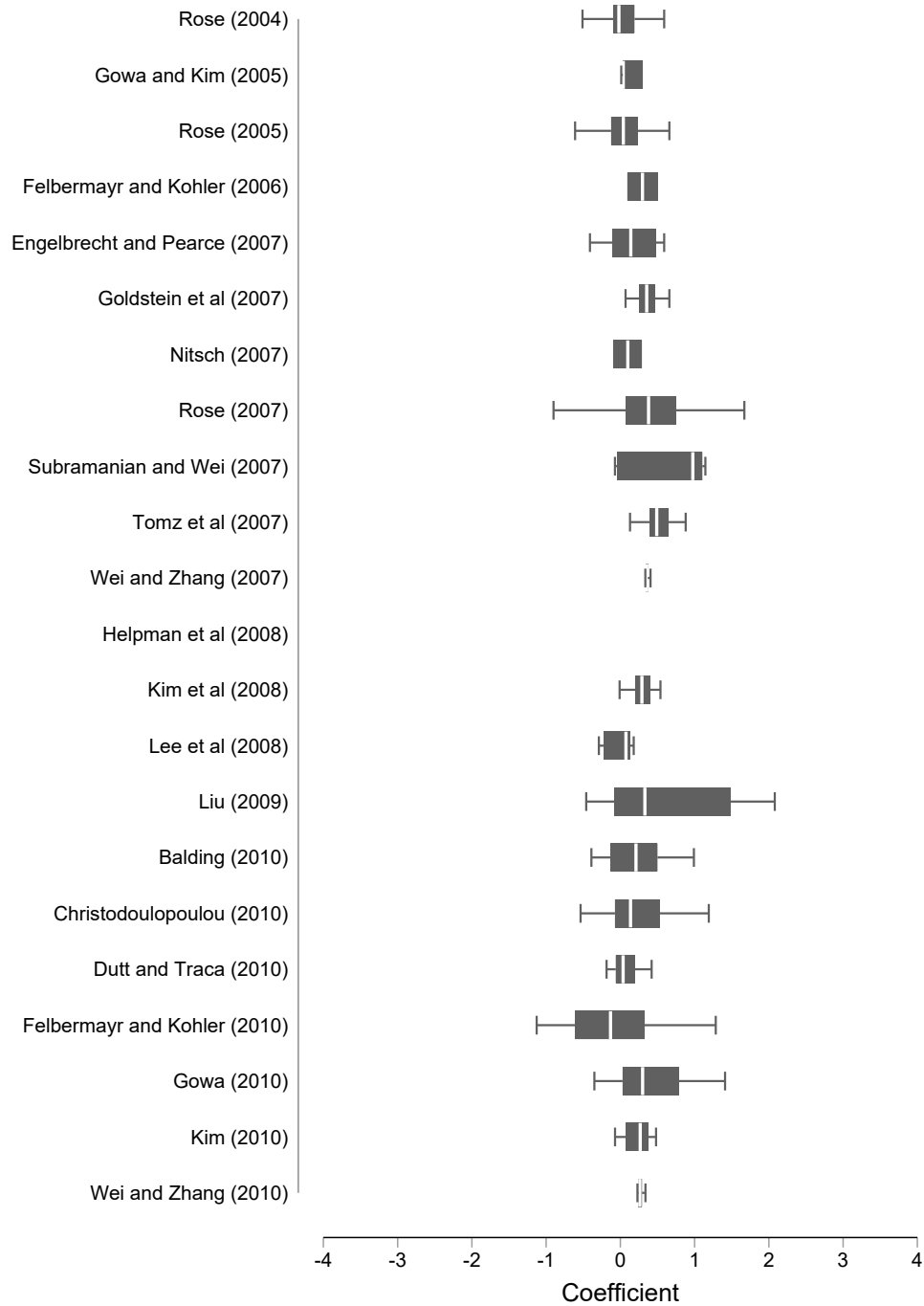


Figure A.3: Box plots for estimates for "Both in WTO" (2004–2010)

**Note:** Papers listed in chronological order. Outside values are not shown. Papers with too few observations to calculate percentiles are not removed from the list.

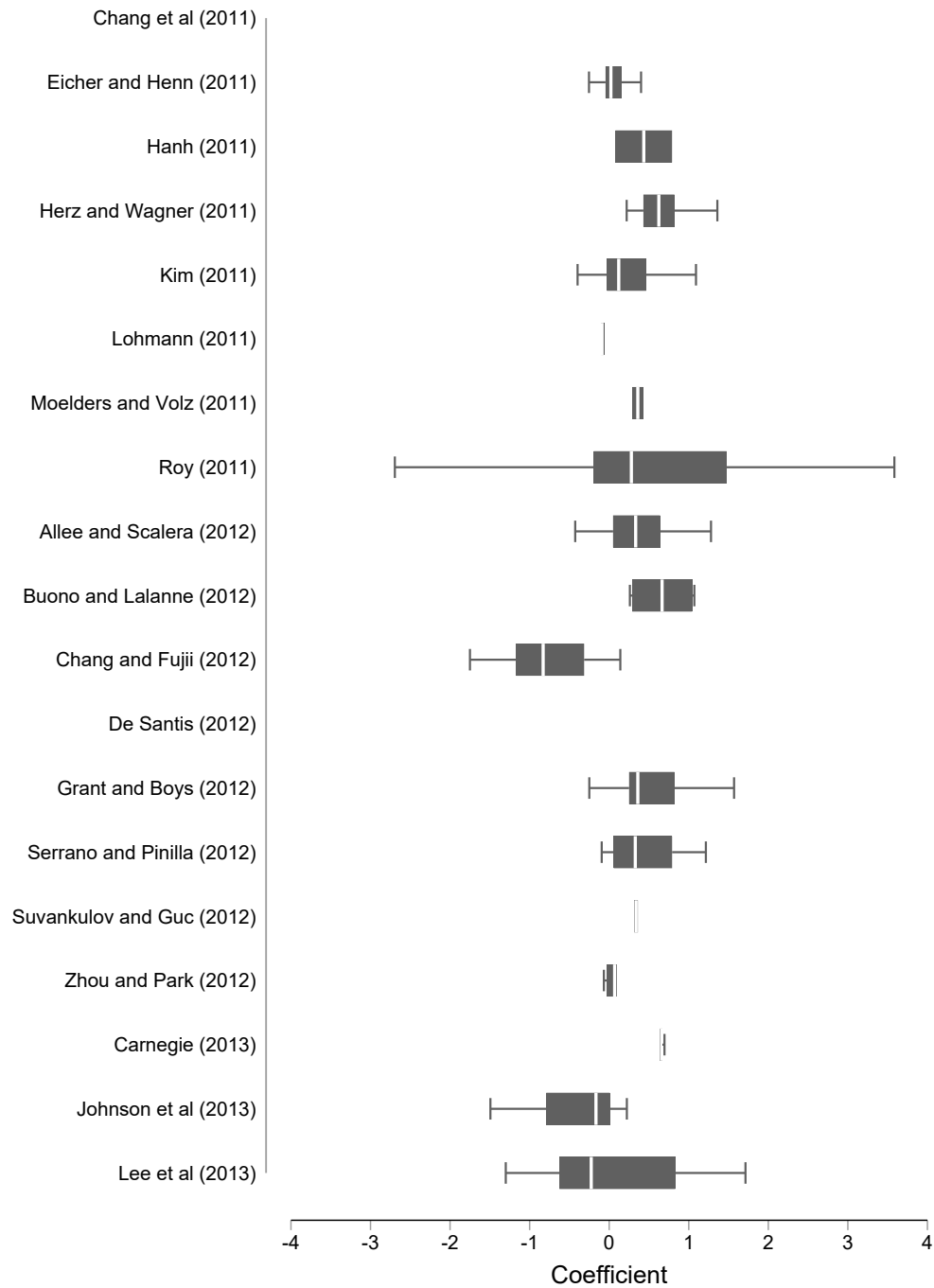


Figure A.4: Box plots for estimates for "Both in WTO" (2010–2013)

**Note:** Papers listed in chronological order. Outside values are not shown. Papers with too few observations to calculate percentiles are not removed from the list.

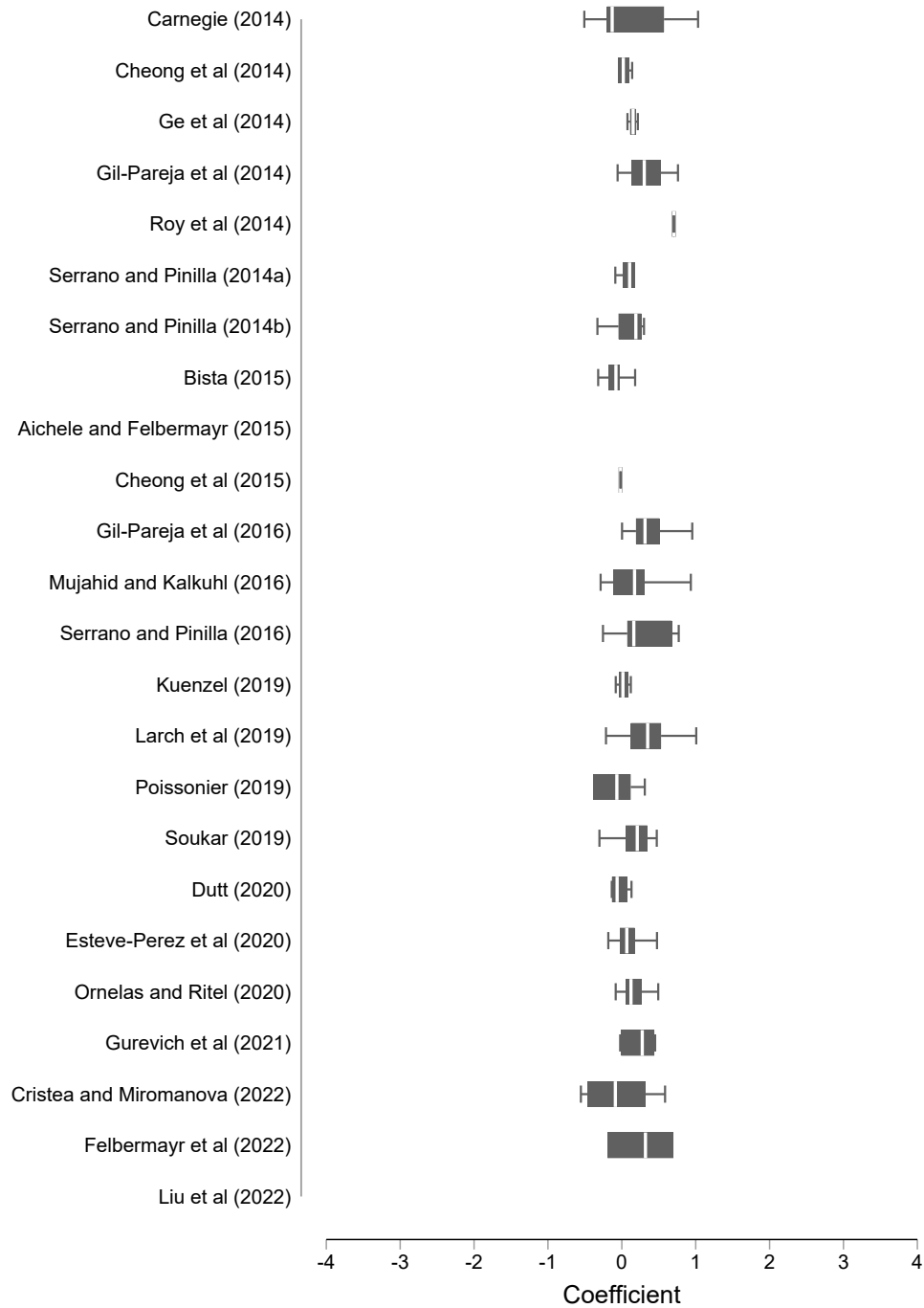


Figure A.5: Box plots for estimates for "Both in WTO" (2014–2022)

**Note:** Papers listed in chronological order. Outside values are not shown. Papers with too few observations to calculate percentiles are not removed from the list.

### A.3 Robustness tests

Table A.1: Average estimates in the literature

|                                           | Mean | 95% CI      | Weighted mean | 95% CI      |
|-------------------------------------------|------|-------------|---------------|-------------|
| <i>Bothin</i> ( $\gamma_1$ ): Both in WTO | 0.21 | [0.18 0.24] | 0.20          | [0.18 0.23] |
| <i>Onein</i> ( $\gamma_2$ ): One in WTO   | 0.13 | [0.10 0.16] | 0.12          | [0.08 0.17] |
| <i>All</i>                                | 0.18 | [0.16 0.20] | 0.17          | [0.14 0.20] |

**Note:** Confidence intervals are calculated without clustering standard errors at the study level.

Table A.2: Tests for publication bias

| VARIABLES    | (1)<br><i>Bothin</i><br>All | (2)<br><i>Bothin</i><br>Journals only | (3)<br><i>Bothin</i><br>FE | (4)<br><i>Onein</i><br>All | (5)<br><i>Onein</i><br>Journals only | (6)<br><i>Onein</i><br>FE |
|--------------|-----------------------------|---------------------------------------|----------------------------|----------------------------|--------------------------------------|---------------------------|
| se           | -0.314<br>(0.211)           | -0.311<br>(0.214)                     | -0.435<br>(0.340)          | 0.002<br>(0.291)           | 0.002<br>(0.277)                     | 0.002<br>(0.352)          |
| Constant     | 0.258***<br>(0.025)         | 0.257***<br>(0.025)                   | 0.275***<br>(0.058)        | 0.138***<br>(0.040)        | 0.159***<br>(0.037)                  | 0.138**<br>(0.062)        |
| Observations | 1,422                       | 1,381                                 | 1,422                      | 862                        | 807                                  | 862                       |
| R-squared    | 0.019                       | 0.019                                 | 0.029                      | 0.013                      | 0.013                                | 0.016                     |
| Number of id |                             |                                       | 62                         |                            |                                      | 34                        |

Standard errors in parentheses  
\*\*\* p<0.01, \*\* p<0.05, \* p<0.1

**Note:** The dependent variable in columns (1)–(3) is the estimate for “*Bothin*”. Column (1) includes all estimates of this type. Column (2) only includes those that were published in a journal. Column (3) includes fixed effects by study. The value reported for the constant is the average of all fixed effects. The dependent variable in columns (4)–(6) is the estimate for “*Onein*”. Column (4) includes all estimates of this type. Column (5) only those that were published in a journal. Column (6) includes fixed effects by study. The value reported for the constant is the average of all fixed effects. Standard errors (in parentheses) are not clustered at the study level and bootstrapped with 10,000 replications.

Table A.3: Tests for publication bias

| VARIABLES      | (1)<br><i>Bothin</i><br>All | (2)<br><i>Bothin</i><br>Journals only | (3)<br><i>Bothin</i><br>FE | (4)<br><i>Onein</i><br>All | (5)<br><i>Onein</i><br>Journals only | (6)<br><i>Onein</i><br>FE |
|----------------|-----------------------------|---------------------------------------|----------------------------|----------------------------|--------------------------------------|---------------------------|
| Standard error | -0.251<br>(0.204)           | -0.245<br>(0.211)                     | -0.339<br>(0.391)          | -0.106<br>(0.255)          | -0.033<br>(0.237)                    | -0.039<br>(0.301)         |
| Constant       | 0.249***<br>(0.040)         | 0.247***<br>(0.041)                   | 0.260***<br>(0.062)        | 0.160***<br>(0.045)        | 0.172***<br>(0.045)                  | 0.151***<br>(0.052)       |
| Observations   | 1,422                       | 1,381                                 | 1,422                      | 862                        | 807                                  | 862                       |
| R-squared      | 0.007                       | 0.007                                 | 0.009                      | 0.002                      | 0.000                                | 0.000                     |
| Number of id   |                             |                                       | 62                         |                            |                                      | 34                        |

Standard errors in parentheses  
\*\*\* p<0.01, \*\* p<0.05, \* p<0.1

**Note:** The dependent variable in columns (1)–(3) is the estimate winsorized at 1% for “*Bothin*”. Column (1) includes all estimates of this type. Column (2) includes only those that were published in a journal. Column (3) includes fixed effects by study. The value reported for the constant is the average of all fixed effects. The dependent variable in columns (4)–(6) is the estimate winsorized at 1% for “*Onein*”. Column (4) includes all estimates of this type. Column (5) only those that were published in a journal. Column (6) includes fixed effects by study. The value reported for the constant is the average of all fixed effects. Standard errors (in parentheses) are clustered at the study level and bootstrapped with 10,000 replications.

Table A.4: Tests for publication bias with sub-samples

| VARIABLES      | (1)<br><i>Bothin</i><br>MTRs + pair FEs | (2)<br><i>Bothin</i><br>no MTRs + no FEs | (3)<br><i>Onein</i><br>MTRs + pair FEs | (4)<br><i>Onein</i><br>no MTRs + no FEs |
|----------------|-----------------------------------------|------------------------------------------|----------------------------------------|-----------------------------------------|
| Standard error | -0.193<br>(0.476)                       | -0.145<br>(0.360)                        | -0.626<br>(0.842)                      | 0.008<br>(0.372)                        |
| Constant       | 0.134**<br>(0.064)                      | 0.192***<br>(0.052)                      | 0.435<br>(0.270)                       | 0.065<br>(0.050)                        |
| Observations   | 317                                     | 629                                      | 33                                     | 527                                     |
| R-squared      | 0.012                                   | 0.001                                    | 0.053                                  | 0.000                                   |

Standard errors in parentheses  
\*\*\* p<0.01, \*\* p<0.05, \* p<0.1

**Note:** The dependent variable in columns (1) and (2) is the estimate for “*Bothin*”. The dependent variable in columns (3) and (4) is the estimate for “*Onein*”. Columns (1) and (3) are based on those studies that include theoretical multilateral resistances and pair fixed effects. Columns (2) and (4) are based on those studies that include neither theoretical multilateral resistances nor pair fixed effects. Standard errors (in parentheses) are clustered at the study level and bootstrapped with 10,000 replications.

Table A.5: Nonlinear model: WAAP

| VARIABLES    | (1)<br>All          | (2)<br><i>Bothin</i> | (3)<br><i>Onein</i> |
|--------------|---------------------|----------------------|---------------------|
| Precision    | 0.317***<br>(0.007) | 0.332***<br>(0.014)  | 0.310***<br>(0.007) |
| Observations | 1,520               | 934                  | 595                 |
| R-squared    | 0.566               | 0.372                | 0.750               |

Standard errors in parentheses

\*\*\* p&lt;0.01, \*\* p&lt;0.05, \* p&lt;0.1

**Note:** WAAP is computed as an unrestricted weighted average of the adequately powered. This model is based on the funnel plot. It discards estimates with retrospective power below 80% and computes an inverse-variance-weighted mean of the remaining estimates. This is used to reduce the publication bias in meta-analysis. The dependent variable in column (1) is the estimate for “*All*”. In column (2) the dependent variable is the estimate for “*Bothin*”. In Column (3) the dependent variable is the estimate for “*Onein*”.

Table A.6: Selection model: Andrews-Kasy

|                | $\mu$ | $\tau$ | <i>P1</i> | <i>P2</i> | <i>P3</i> |
|----------------|-------|--------|-----------|-----------|-----------|
| Estimate       | 0.049 | 0.448  | 0.312     | 0.894     | 0.839     |
| Standard error | 0.025 | 0.016  | 0.047     | 0.085     | 0.067     |

**Note:** Estimates are derived from the Andrews and Kasy (2019) selection model, which accounts for publication bias by comparing the likelihood of publishing results across different significance levels and signs of the estimated effect. The parameter  $\mu$  represents the publication-bias-adjusted mean effect size across studies, while  $\tau$  denotes the standard deviation of the effect. P1 indicates the probability of publishing significant negative results relative to the overall probability of publishing significant estimates. P2 refers to the probability of publishing negative and statistically insignificant results relative to the probability of publishing significant estimates. P3 captures the probability of publishing positive but statistically insignificant results relative to the probability of publishing significant estimates.

Table A.7: MAIVE: Meta-Analysis Instrumental Variable Estimator

|                            |                 |
|----------------------------|-----------------|
| MAIVE coefficient          | 0.082           |
| MAIVE standard error       | 0.238           |
| F-test of first step in IV | 0.474           |
| Hausman-type test          | 0.966           |
| Critical Value of Chi2(1)  | 3.841           |
| AR Confidence interval     | [−0.781, 0.999] |

**Note:** The dependent variable is the estimate for “*All*” WTO effect. MAIVE takes the inverse of the sample size of primary studies as an instrument for reported squared standard errors. Standard errors in parentheses are clustered at the study level. The Hausman test statistic consists of a weighted squared difference between the MAIVE and a standard point estimate, which is the same method as MAIVE but without instrumenting the standard errors and including inverse-variance weights.

Table A.8: Estimation characteristics (*Bothin*)

| VARIABLES               | (1)                 | (2)                 | (3)                 | (4)                 | (5)                  | (6)                 | (7)                  |
|-------------------------|---------------------|---------------------|---------------------|---------------------|----------------------|---------------------|----------------------|
| Standard error          | -0.251<br>(0.204)   | -0.229<br>(0.206)   | -0.235<br>(0.199)   | -0.236<br>(0.206)   | -0.251<br>(0.204)    | -0.241<br>(0.196)   | -0.192<br>(0.219)    |
| No or atheoretical MTRs |                     | 0.075<br>(0.076)    |                     |                     |                      |                     | 0.054<br>(0.072)     |
| Other than PPML         |                     |                     | 0.236***<br>(0.066) |                     |                      |                     | 0.240***<br>(0.069)  |
| No pair fe              |                     |                     |                     | -0.063<br>(0.055)   |                      |                     | -0.096**<br>(0.046)  |
| No domestic trade       |                     |                     |                     |                     | -0.168***<br>(0.052) |                     | -0.350***<br>(0.072) |
| One advanced            |                     |                     |                     |                     |                      | -0.100<br>(0.109)   | -0.008<br>(0.078)    |
| Both advanced           |                     |                     |                     |                     |                      | 0.066<br>(0.106)    | 0.058<br>(0.089)     |
| Constant                | 0.249***<br>(0.040) | 0.197***<br>(0.070) | 0.071<br>(0.062)    | 0.282***<br>(0.051) | 0.414***<br>(0.042)  | 0.251***<br>(0.037) | 0.420***<br>(0.055)  |
| Observations            | 1,422               | 1,422               | 1,422               | 1,422               | 1,422                | 1,422               | 1,422                |
| R-squared               | 0.007               | 0.013               | 0.057               | 0.012               | 0.009                | 0.013               | 0.075                |

Standard errors in parentheses  
\*\*\* p<0.01, \*\* p<0.05, \* p<0.1

**Note:** The dependent variable is the estimate winsorized at 1% for “*Bothin*”. Standard errors (in parentheses) are clustered at the study level and bootstrapped with 10,000 replications.

Table A.9: Estimation characteristics (*Onein*)

| VARIABLES               | (1)                 | (2)                | (3)                | (4)                 | (5)                  | (6)                 | (7)                  |
|-------------------------|---------------------|--------------------|--------------------|---------------------|----------------------|---------------------|----------------------|
| Standard error          | -0.106<br>(0.255)   | -0.205<br>(0.322)  | -0.099<br>(0.254)  | -0.076<br>(0.277)   | -0.112<br>(0.260)    | -0.188<br>(0.227)   | -0.197<br>(0.268)    |
| No or atheoretical MTRs |                     | -0.201<br>(0.178)  |                    |                     |                      |                     | -0.067<br>(0.248)    |
| Other than PPML         |                     |                    | -0.197<br>(0.149)  |                     |                      |                     | -0.030<br>(0.105)    |
| No pair FE              |                     |                    |                    | -0.157**<br>(0.070) |                      |                     | -0.198***<br>(0.062) |
| No domestic trade       |                     |                    |                    |                     | -0.482***<br>(0.049) |                     | -0.275<br>(0.284)    |
| One advanced            |                     |                    |                    |                     |                      | 0.353**<br>(0.173)  | 0.409*<br>(0.222)    |
| Both advanced           |                     |                    |                    |                     |                      | 0.271**<br>(0.116)  | 0.294**<br>(0.124)   |
| Constant                | 0.160***<br>(0.045) | 0.356**<br>(0.169) | 0.348**<br>(0.143) | 0.261***<br>(0.044) | 0.636***<br>(0.042)  | 0.123***<br>(0.044) | 0.609***<br>(0.049)  |
| Observations            | 862                 | 862                | 862                | 862                 | 862                  | 862                 | 862                  |
| R-squared               | 0.002               | 0.016              | 0.008              | 0.027               | 0.015                | 0.063               | 0.118                |

Standard errors in parentheses  
\*\*\* p<0.01, \*\* p<0.05, \* p<0.1

**Note:** The dependent variable is the estimate winsorized at 1% for “*Onein*”. Standard errors (in parentheses) are clustered at the study level and bootstrapped with 10,000 replications.

## B Gravity results: Tables and other trade databases

Table B.1: Deviations from standard practices

|                               | Both in      | Upper bound<br>TA | EU          | Both in      | Lower bound<br>TA | EU          |
|-------------------------------|--------------|-------------------|-------------|--------------|-------------------|-------------|
| <i>IMF DOTS (importer)</i>    |              |                   |             |              |                   |             |
| Theory-consistent             | 0.43 (0.06)  | 0.20 (0.06)       | 0.67 (0.07) | 0.32 (0.06)  | 0.13 (0.05)       | 0.55 (0.07) |
| OLS on log(trade)             | 0.14 (0.03)  | 0.30 (0.03)       | 0.97 (0.05) | 0.16 (0.03)  | 0.30 (0.03)       | 0.97 (0.05) |
| Non-theory MRT                | 0.40 (0.06)  | 0.18 (0.05)       | 0.53 (0.06) | 0.28 (0.05)  | 0.12 (0.05)       | 0.43 (0.06) |
| No pair FE                    | 0.70 (0.06)  | 0.26 (0.08)       | 0.14 (0.10) | 0.66 (0.06)  | 0.22 (0.08)       | 0.15 (0.10) |
| No domestic trade             | -0.05 (0.07) | 0.02 (0.05)       | 0.47 (0.04) | -0.05 (0.07) | 0.02 (0.05)       | 0.47 (0.04) |
| <i>IMF DOTS (exporter)</i>    |              |                   |             |              |                   |             |
| Theory-consistent             | 0.47 (0.06)  | 0.28 (0.05)       | 0.69 (0.07) | 0.34 (0.06)  | 0.18 (0.04)       | 0.53 (0.07) |
| OLS on log(trade)             | 0.14 (0.03)  | 0.36 (0.03)       | 0.68 (0.05) | 0.17 (0.03)  | 0.37 (0.03)       | 0.68 (0.05) |
| Non-theory MRT                | 0.41 (0.05)  | 0.25 (0.05)       | 0.54 (0.05) | 0.28 (0.05)  | 0.16 (0.04)       | 0.41 (0.06) |
| No pair FE                    | 0.65 (0.07)  | 0.28 (0.08)       | 0.20 (0.10) | 0.60 (0.07)  | 0.24 (0.07)       | 0.21 (0.10) |
| No domestic trade             | -0.11 (0.07) | 0.07 (0.04)       | 0.39 (0.04) | -0.11 (0.07) | 0.07 (0.04)       | 0.39 (0.04) |
| <i>UN COMTRADE (importer)</i> |              |                   |             |              |                   |             |
| Theory-consistent             | 0.41 (0.06)  | 0.18 (0.06)       | 0.55 (0.05) | 0.32 (0.06)  | 0.13 (0.05)       | 0.43 (0.06) |
| OLS on log(trade)             | 0.11 (0.03)  | 0.21 (0.03)       | 0.55 (0.04) | 0.09 (0.03)  | 0.21 (0.03)       | 0.55 (0.04) |
| Non-theory MRT                | 0.39 (0.06)  | 0.17 (0.05)       | 0.45 (0.05) | 0.29 (0.06)  | 0.12 (0.05)       | 0.36 (0.05) |
| No pair FE                    | 0.56 (0.06)  | 0.20 (0.08)       | 0.08 (0.10) | 0.55 (0.07)  | 0.17 (0.08)       | 0.09 (0.09) |
| No domestic trade             | -0.06 (0.07) | -0.00 (0.05)      | 0.31 (0.04) | -0.06 (0.07) | -0.00 (0.05)      | 0.31 (0.04) |
| <i>UN COMTRADE (exporter)</i> |              |                   |             |              |                   |             |
| Theory-consistent             | 0.40 (0.06)  | 0.27 (0.05)       | 0.60 (0.06) | 0.30 (0.06)  | 0.20 (0.04)       | 0.47 (0.06) |
| OLS on log(trade)             | 0.12 (0.04)  | 0.23 (0.02)       | 0.53 (0.04) | 0.10 (0.04)  | 0.22 (0.02)       | 0.53 (0.04) |
| Non-theory MRT                | 0.38 (0.06)  | 0.24 (0.05)       | 0.49 (0.05) | 0.28 (0.05)  | 0.18 (0.04)       | 0.39 (0.05) |
| No pair FE                    | 0.52 (0.07)  | 0.27 (0.08)       | 0.16 (0.10) | 0.51 (0.07)  | 0.25 (0.08)       | 0.17 (0.10) |
| No domestic trade             | -0.40 (0.14) | 0.07 (0.04)       | 0.31 (0.04) | -0.40 (0.14) | 0.07 (0.04)       | 0.31 (0.04) |
| <i>BACI CEPII</i>             |              |                   |             |              |                   |             |
| Theory-consistent             | 0.25 (0.05)  | 0.07 (0.04)       | 0.60 (0.05) | 0.20 (0.05)  | 0.07 (0.03)       | 0.53 (0.05) |
| OLS on log(trade)             | 0.11 (0.05)  | 0.03 (0.02)       | 0.33 (0.04) | 0.10 (0.05)  | 0.02 (0.02)       | 0.33 (0.04) |
| Non-theory MRT                | 0.20 (0.04)  | 0.04 (0.03)       | 0.41 (0.03) | 0.17 (0.04)  | 0.05 (0.03)       | 0.37 (0.04) |
| No pair FE                    | 0.65 (0.07)  | 0.23 (0.07)       | 0.13 (0.09) | 0.66 (0.07)  | 0.23 (0.07)       | 0.13 (0.09) |
| No domestic trade             | 0.08 (0.06)  | 0.05 (0.03)       | 0.14 (0.04) | 0.08 (0.06)  | 0.05 (0.03)       | 0.14 (0.04) |
| <i>TradeProd CEPII</i>        |              |                   |             |              |                   |             |
| Theory-consistent             | 0.31 (0.06)  | 0.23 (0.08)       | 0.62 (0.06) | 0.15 (0.05)  | 0.05 (0.06)       | 0.41 (0.06) |
| OLS on log(trade)             | 0.18 (0.04)  | 0.19 (0.03)       | 0.48 (0.04) | 0.15 (0.04)  | 0.19 (0.03)       | 0.48 (0.04) |
| Non-theory MRT                | 0.63 (0.16)  | 0.15 (0.12)       | 0.47 (0.10) | 0.41 (0.09)  | 0.06 (0.07)       | 0.39 (0.06) |
| No pair FE                    | 0.97 (0.07)  | 0.39 (0.09)       | 0.20 (0.11) | 0.98 (0.07)  | 0.36 (0.09)       | 0.22 (0.11) |
| No domestic trade             | -0.01 (0.08) | -0.03 (0.07)      | 0.20 (0.04) | -0.01 (0.08) | -0.03 (0.07)      | 0.20 (0.04) |
| <i>WTO Structural DB</i>      |              |                   |             |              |                   |             |
| Theory-consistent             | 0.55 (0.08)  | 0.41 (0.07)       | 0.73 (0.07) | 0.26 (0.06)  | 0.14 (0.04)       | 0.42 (0.06) |
| OLS on log(trade)             | 0.13 (0.04)  | 0.17 (0.02)       | 0.49 (0.04) | 0.10 (0.04)  | 0.17 (0.02)       | 0.48 (0.04) |
| Non-theory MRT                | 0.77 (0.24)  | 0.50 (0.20)       | 0.63 (0.15) | 0.61 (0.12)  | 0.26 (0.06)       | 0.35 (0.07) |
| No pair FE                    | 1.35 (0.10)  | 0.58 (0.09)       | 0.32 (0.12) | 1.24 (0.09)  | 0.53 (0.09)       | 0.36 (0.12) |
| No domestic trade             | 0.02 (0.07)  | 0.03 (0.05)       | 0.17 (0.04) | 0.02 (0.07)  | 0.03 (0.05)       | 0.17 (0.04) |

**Note:** The table reports the point estimates and standard errors (clustered by country pair) for *Bothin*, *TA*, and *EU*. “Upper bound”: specifications do not control for border-year dummies. “Lower bound”: specifications control for border-year dummies. The rows correspond to the different specifications used in the estimation.

Table B.2: Robustness

|                               | Both in     | Upper bound<br>TA | EU          | Both in     | Lower bound<br>TA | EU          |
|-------------------------------|-------------|-------------------|-------------|-------------|-------------------|-------------|
| <i>IMF DOTS (importer)</i>    |             |                   |             |             |                   |             |
| Theory-consistent             | 0.43 (0.06) | 0.20 (0.06)       | 0.67 (0.07) | 0.32 (0.06) | 0.13 (0.05)       | 0.55 (0.07) |
| Distance x year               | 0.36 (0.05) | 0.17 (0.05)       | 0.59 (0.07) | 0.34 (0.05) | 0.11 (0.04)       | 0.52 (0.08) |
| 3-way clustering              | 0.43 (0.12) | 0.20 (0.07)       | 0.67 (0.11) | 0.32 (0.11) | 0.13 (0.06)       | 0.55 (0.11) |
| <i>IMF DOTS (exporter)</i>    |             |                   |             |             |                   |             |
| Theory-consistent             | 0.47 (0.06) | 0.28 (0.05)       | 0.69 (0.07) | 0.34 (0.06) | 0.18 (0.04)       | 0.53 (0.07) |
| Distance x year               | 0.37 (0.05) | 0.23 (0.04)       | 0.59 (0.07) | 0.35 (0.05) | 0.14 (0.04)       | 0.52 (0.07) |
| 3-way clustering              | 0.47 (0.13) | 0.28 (0.06)       | 0.69 (0.11) | 0.34 (0.12) | 0.18 (0.04)       | 0.53 (0.11) |
| <i>UN COMTRADE (importer)</i> |             |                   |             |             |                   |             |
| Theory-consistent             | 0.41 (0.06) | 0.18 (0.06)       | 0.55 (0.05) | 0.32 (0.06) | 0.13 (0.05)       | 0.43 (0.06) |
| Distance x year               | 0.34 (0.05) | 0.16 (0.05)       | 0.47 (0.05) | 0.33 (0.05) | 0.12 (0.04)       | 0.43 (0.06) |
| 3-way clustering              | 0.41 (0.11) | 0.18 (0.07)       | 0.55 (0.10) | 0.32 (0.10) | 0.13 (0.06)       | 0.43 (0.10) |
| <i>UN COMTRADE (exporter)</i> |             |                   |             |             |                   |             |
| Theory-consistent             | 0.40 (0.06) | 0.27 (0.05)       | 0.60 (0.06) | 0.30 (0.06) | 0.20 (0.04)       | 0.47 (0.06) |
| Distance x year               | 0.32 (0.06) | 0.23 (0.04)       | 0.52 (0.06) | 0.30 (0.06) | 0.17 (0.03)       | 0.47 (0.06) |
| 3-way clustering              | 0.40 (0.13) | 0.27 (0.06)       | 0.60 (0.11) | 0.30 (0.12) | 0.20 (0.04)       | 0.47 (0.11) |
| <i>BACI CEPII</i>             |             |                   |             |             |                   |             |
| Theory-consistent             | 0.25 (0.05) | 0.07 (0.04)       | 0.60 (0.05) | 0.20 (0.05) | 0.07 (0.03)       | 0.53 (0.05) |
| Distance x year               | 0.20 (0.05) | 0.06 (0.04)       | 0.55 (0.05) | 0.20 (0.05) | 0.06 (0.04)       | 0.54 (0.05) |
| 3-way clustering              | 0.25 (0.13) | 0.07 (0.03)       | 0.60 (0.10) | 0.20 (0.12) | 0.07 (0.01)       | 0.53 (0.10) |
| <i>TradeProd CEPII</i>        |             |                   |             |             |                   |             |
| Theory-consistent             | 0.31 (0.06) | 0.23 (0.08)       | 0.62 (0.06) | 0.15 (0.05) | 0.05 (0.06)       | 0.41 (0.06) |
| Distance x year               | 0.17 (0.04) | 0.12 (0.07)       | 0.46 (0.06) | 0.16 (0.04) | 0.04 (0.05)       | 0.41 (0.06) |
| 3-way clustering              | 0.31 (0.08) | 0.23 (0.11)       | 0.62 (0.10) | 0.15 (0.06) | 0.05 (0.07)       | 0.41 (0.11) |
| <i>WTO Structural DB</i>      |             |                   |             |             |                   |             |
| Theory-consistent             | 0.55 (0.08) | 0.41 (0.07)       | 0.73 (0.07) | 0.26 (0.06) | 0.14 (0.04)       | 0.42 (0.06) |
| Distance x year               | 0.34 (0.06) | 0.24 (0.06)       | 0.52 (0.06) | 0.26 (0.05) | 0.14 (0.04)       | 0.42 (0.06) |
| 3-way clustering              | 0.55 (0.13) | 0.41 (0.10)       | 0.73 (0.11) | 0.26 (0.10) | 0.14 (0.05)       | 0.42 (0.12) |

**Note:** The table reports the point estimates and standard errors (clustered by country pair) for *Bothin*, *TA*, and *EU*. “Upper bound”: specifications do not control for border-year dummies. “Lower bound”: specifications control for border-year dummies. The rows correspond to the different specifications used in the estimation.

Table B.3: Results by country income level

|                               | Upper bound  | Lower bound  |
|-------------------------------|--------------|--------------|
| <i>IMF DOTS (importer)</i>    |              |              |
| AEs-AEs                       | 0.08 (0.13)  | -0.07 (0.13) |
| AEs-EMDEs                     | 0.42 (0.09)  | 0.28 (0.09)  |
| EMDEs-AEs                     | 0.38 (0.08)  | 0.27 (0.08)  |
| EMDEs-EMDEs                   | 0.60 (0.09)  | 0.52 (0.09)  |
| TA                            | 0.20 (0.06)  | 0.13 (0.05)  |
| EU                            | 0.68 (0.07)  | 0.55 (0.07)  |
| <i>IMF DOTS (exporter)</i>    |              |              |
| AEs-AEs                       | 0.33 (0.10)  | 0.18 (0.11)  |
| AEs-EMDEs                     | 0.45 (0.09)  | 0.30 (0.08)  |
| EMDEs-AEs                     | 0.45 (0.09)  | 0.32 (0.08)  |
| EMDEs-EMDEs                   | 0.55 (0.08)  | 0.44 (0.08)  |
| TA                            | 0.28 (0.05)  | 0.18 (0.04)  |
| EU                            | 0.69 (0.07)  | 0.53 (0.07)  |
| <i>UN COMTRADE (importer)</i> |              |              |
| AEs-AEs                       | -0.07 (0.13) | -0.20 (0.13) |
| AEs-EMDEs                     | 0.37 (0.10)  | 0.27 (0.09)  |
| EMDEs-AEs                     | 0.35 (0.08)  | 0.27 (0.08)  |
| EMDEs-EMDEs                   | 0.57 (0.09)  | 0.52 (0.09)  |
| TA                            | 0.19 (0.06)  | 0.13 (0.05)  |
| EU                            | 0.55 (0.05)  | 0.44 (0.06)  |
| <i>UN COMTRADE (exporter)</i> |              |              |
| AEs-AEs                       | 0.23 (0.11)  | 0.12 (0.12)  |
| AEs-EMDEs                     | 0.44 (0.09)  | 0.35 (0.08)  |
| EMDEs-AEs                     | 0.33 (0.10)  | 0.24 (0.09)  |
| EMDEs-EMDEs                   | 0.50 (0.09)  | 0.44 (0.08)  |
| TA                            | 0.27 (0.05)  | 0.19 (0.04)  |
| EU                            | 0.60 (0.06)  | 0.47 (0.06)  |
| <i>BACI CEPII</i>             |              |              |
| AEs-AEs                       | 0.04 (0.09)  | -0.03 (0.09) |
| AEs-EMDEs                     | 0.27 (0.08)  | 0.21 (0.07)  |
| EMDEs-AEs                     | 0.22 (0.06)  | 0.18 (0.06)  |
| EMDEs-EMDEs                   | 0.31 (0.08)  | 0.29 (0.08)  |
| TA                            | 0.07 (0.04)  | 0.07 (0.03)  |
| EU                            | 0.61 (0.05)  | 0.53 (0.05)  |
| <i>TradeProd CEPII</i>        |              |              |
| AEs-AEs                       | 0.59 (0.18)  | 0.32 (0.14)  |
| AEs-EMDEs                     | 0.39 (0.07)  | 0.22 (0.06)  |
| EMDEs-AEs                     | 0.14 (0.10)  | -0.04 (0.09) |
| EMDEs-EMDEs                   | 0.41 (0.08)  | 0.31 (0.07)  |
| TA                            | 0.23 (0.08)  | 0.05 (0.05)  |
| EU                            | 0.62 (0.05)  | 0.41 (0.06)  |
| <i>WTO Structural DB</i>      |              |              |
| AEs-AEs                       | 0.37 (0.08)  | 0.08 (0.06)  |
| AEs-EMDEs                     | 0.62 (0.08)  | 0.30 (0.07)  |
| EMDEs-AEs                     | 0.49 (0.12)  | 0.18 (0.09)  |
| EMDEs-EMDEs                   | 0.55 (0.09)  | 0.33 (0.08)  |
| TA                            | 0.41 (0.07)  | 0.14 (0.04)  |
| EU                            | 0.73 (0.07)  | 0.42 (0.06)  |

**Note:** The table reports the point estimates and standard errors (clustered by country pair) for *Bothin*, *TA*, and *EU*. “Upper bound”: specifications do not control for border-year dummies. “Lower bound”: specifications control for border-year dummies. The rows correspond to the different specifications used in the estimation.

Table B.4: Results by country income level in different periods and databases

|                               | Upper bound               |                          |                      | Lower bound               |                          |                      |
|-------------------------------|---------------------------|--------------------------|----------------------|---------------------------|--------------------------|----------------------|
|                               | Early GATT<br>(1948–1973) | Late GATT<br>(1974–1994) | WTO<br>(1995–onward) | Early GATT<br>(1948–1973) | Late GATT<br>(1974–1994) | WTO<br>(1995–onward) |
| <i>IMF DOTS (importer)</i>    |                           |                          |                      |                           |                          |                      |
| AEs-AEs                       | 0.28 (0.05)               | -0.23 (0.18)             | 0.01 (0.09)          | 0.32 (0.07)               | -0.18 (0.20)             | -0.06 (0.09)         |
| AEs-EMDEs                     | -0.28 (0.07)              | 0.10 (0.11)              | 0.32 (0.07)          | -0.30 (0.08)              | 0.15 (0.12)              | 0.23 (0.07)          |
| EMDEs-AEs                     | -0.04 (0.08)              | 0.24 (0.12)              | 0.25 (0.07)          | -0.06 (0.09)              | 0.28 (0.12)              | 0.18 (0.07)          |
| EMDEs-EMDEs                   | -0.24 (0.08)              | 0.23 (0.07)              | 0.36 (0.09)          | -0.28 (0.09)              | 0.27 (0.07)              | 0.32 (0.09)          |
| <i>IMF DOTS (exporter)</i>    |                           |                          |                      |                           |                          |                      |
| AEs-AEs                       | 0.28 (0.05)               | -0.25 (0.28)             | 0.16 (0.10)          | 0.31 (0.07)               | -0.22 (0.29)             | 0.08 (0.09)          |
| AEs-EMDEs                     | -0.21 (0.07)              | 0.02 (0.11)              | 0.31 (0.08)          | -0.24 (0.07)              | 0.06 (0.12)              | 0.21 (0.07)          |
| EMDEs-AEs                     | -0.05 (0.09)              | 0.19 (0.13)              | 0.37 (0.07)          | -0.07 (0.10)              | 0.23 (0.14)              | 0.29 (0.07)          |
| EMDEs-EMDEs                   | -0.19 (0.08)              | 0.18 (0.07)              | 0.29 (0.07)          | -0.23 (0.09)              | 0.21 (0.07)              | 0.23 (0.07)          |
| <i>UN COMTRADE (importer)</i> |                           |                          |                      |                           |                          |                      |
| AEs-AEs                       | 0.24 (0.05)               | -0.09 (0.09)             | 0.03 (0.09)          | 0.13 (0.06)               | -0.04 (0.09)             | -0.03 (0.09)         |
| AEs-EMDEs                     | -0.09 (0.06)              | 0.05 (0.10)              | 0.31 (0.07)          | -0.22 (0.07)              | 0.09 (0.10)              | 0.26 (0.07)          |
| EMDEs-AEs                     | -0.03 (0.05)              | 0.26 (0.10)              | 0.27 (0.07)          | -0.11 (0.06)              | 0.30 (0.10)              | 0.23 (0.07)          |
| EMDEs-EMDEs                   | 0.02 (0.07)               | 0.22 (0.08)              | 0.37 (0.09)          | -0.08 (0.08)              | 0.25 (0.08)              | 0.36 (0.09)          |
| <i>UN COMTRADE (exporter)</i> |                           |                          |                      |                           |                          |                      |
| AEs-AEs                       | 0.24 (0.05)               | -0.16 (0.27)             | 0.18 (0.09)          | 0.14 (0.06)               | -0.12 (0.28)             | 0.12 (0.09)          |
| AEs-EMDEs                     | -0.03 (0.06)              | 0.03 (0.10)              | 0.38 (0.08)          | -0.12 (0.07)              | 0.07 (0.10)              | 0.32 (0.07)          |
| EMDEs-AEs                     | -0.09 (0.08)              | 0.22 (0.12)              | 0.29 (0.07)          | -0.20 (0.08)              | 0.26 (0.13)              | 0.25 (0.07)          |
| EMDEs-EMDEs                   | 0.06 (0.09)               | 0.20 (0.08)              | 0.31 (0.08)          | -0.04 (0.09)              | 0.24 (0.08)              | 0.29 (0.08)          |
| <i>BACI CEPII</i>             |                           |                          |                      |                           |                          |                      |
| AEs-AEs                       | —                         | —                        | 0.38 (0.10)          | —                         | —                        | 0.31 (0.10)          |
| AEs-EMDEs                     | —                         | —                        | -0.05 (0.12)         | —                         | —                        | -0.14 (0.11)         |
| EMDEs-AEs                     | —                         | —                        | 0.04 (0.08)          | —                         | —                        | -0.03 (0.08)         |
| EMDEs-EMDEs                   | —                         | —                        | -0.12 (0.10)         | —                         | —                        | -0.18 (0.10)         |
| <i>TradeProd CEPII</i>        |                           |                          |                      |                           |                          |                      |
| AEs-AEs                       | 0.38 (0.10)               | 0.54 (0.12)              | -0.09 (0.08)         | 0.31 (0.10)               | 0.37 (0.12)              | -0.16 (0.08)         |
| AEs-EMDEs                     | -0.05 (0.12)              | 0.19 (0.08)              | 0.18 (0.05)          | -0.14 (0.11)              | 0.03 (0.08)              | 0.09 (0.05)          |
| EMDEs-AEs                     | 0.04 (0.08)               | 0.04 (0.09)              | 0.01 (0.08)          | -0.03 (0.08)              | -0.15 (0.09)             | -0.08 (0.08)         |
| EMDEs-EMDEs                   | -0.12 (0.10)              | -0.04 (0.15)             | 0.25 (0.08)          | -0.18 (0.10)              | -0.19 (0.15)             | 0.20 (0.07)          |
| <i>WTO Structural DB</i>      |                           |                          |                      |                           |                          |                      |
| AEs-AEs                       | —                         | 0.08 (0.17)              | 0.15 (0.10)          | —                         | 0.06 (0.09)              | 0.15 (0.10)          |
| AEs-EMDEs                     | —                         | 0.27 (0.09)              | 0.33 (0.06)          | —                         | 0.15 (0.05)              | 0.33 (0.06)          |
| EMDEs-AEs                     | —                         | 0.21 (0.12)              | 0.29 (0.07)          | —                         | 0.12 (0.07)              | 0.29 (0.07)          |
| EMDEs-EMDEs                   | —                         | 0.43 (0.10)              | 0.29 (0.07)          | —                         | 0.15 (0.07)              | 0.29 (0.07)          |

**Note:** Results in the first three columns do not control for border-year dummies (“upper bound”). Results in the last three columns control for border-year dummies (“lower bound”). The rows correspond to the different specifications used in the estimation. IMF DOTS reports bilateral trade data from 1948 onwards. UN COMTRADE reports bilateral trade from 1962 onwards. BACII reports bilateral trade after 1996. TradeProd CEPII reports bilateral trade from 1966 onwards. WTO Structural DB reports bilateral trade from 1980 onwards.

Table B.5: GATT and WTO estimated separately

|                        | Upper bound |             | Lower bound |             |
|------------------------|-------------|-------------|-------------|-------------|
|                        | GATT        | WTO         | GATT        | WTO         |
| IMF DOTS (importer)    | 0.27 (0.12) | 0.27 (0.03) | 0.28 (0.13) | 0.22 (0.04) |
| IMF DOTS (exporter)    | 0.26 (0.13) | 0.31 (0.04) | 0.26 (0.14) | 0.20 (0.04) |
| UN COMTRADE (importer) | 0.30 (0.13) | 0.26 (0.03) | 0.34 (0.14) | 0.28 (0.04) |
| UN COMTRADE (exporter) | 0.34 (0.15) | 0.27 (0.03) | 0.38 (0.16) | 0.25 (0.04) |
| BACI CEPII             | —           | 0.25 (0.05) | —           | 0.20 (0.05) |
| TradeProd CEPII        | 0.48 (0.08) | 0.40 (0.03) | 0.35 (0.07) | 0.09 (0.04) |
| WTO Structural DB      | 0.15 (0.09) | 0.62 (0.03) | 0.03 (0.09) | 0.12 (0.04) |

Table B.6: Both in WTO (domestic trade: Head and Mayer, 2021)

|                        | Upper bound | Lower bound |
|------------------------|-------------|-------------|
| IMF DOTS (importer)    | 0.52 (0.06) | 0.20 (0.06) |
| IMF DOTS (exporter)    | 0.50 (0.07) | 0.17 (0.06) |
| UN COMTRADE (importer) | 0.47 (0.06) | 0.17 (0.06) |
| UN COMTRADE (exporter) | 0.41 (0.08) | 0.14 (0.06) |
| BACI CEPII             | 0.22 (0.04) | 0.21 (0.05) |

**Note:** The table reports the point estimates and standard errors for the variable *Bothin*, using domestic trade calculated by Head and Mayer (2021). Standard errors are clustered by country pair. Results in the first two columns do not control for border-year dummies (“upper bound”). Results in the last two columns control for border-year dummies (“lower bound”). The rows correspond to various databases used in the estimation.

Table B.7: Both in WTO versus One in (domestic trade: Head and Mayer, 2021)

|                        | Upper bound |             | Lower bound |             |
|------------------------|-------------|-------------|-------------|-------------|
|                        | Both in     | One in      | Both in     | One in      |
| IMF DOTS (importer)    | 0.76 (0.12) | 0.24 (0.09) | 0.30 (0.11) | 0.10 (0.08) |
| IMF DOTS (exporter)    | 0.76 (0.12) | 0.27 (0.09) | 0.29 (0.10) | 0.12 (0.08) |
| UN COMTRADE (importer) | 0.67 (0.14) | 0.21 (0.11) | 0.28 (0.12) | 0.12 (0.10) |
| UN COMTRADE (exporter) | 0.87 (0.17) | 0.45 (0.15) | 0.50 (0.16) | 0.36 (0.14) |
| BACI CEPII             | 0.30 (0.09) | 0.08 (0.07) | 0.30 (0.11) | 0.09 (0.09) |

**Note:** The table reports the point estimates and standard errors for the variables *Bothin* and *Onein*, using domestic trade calculated by Head and Mayer (2021). Standard errors are clustered by country pair. Results in the first two columns do not control for border-year dummies (“upper bound”). Results in the last two columns control for border-year dummies (“lower bound”). The rows correspond to various databases used in the estimation.

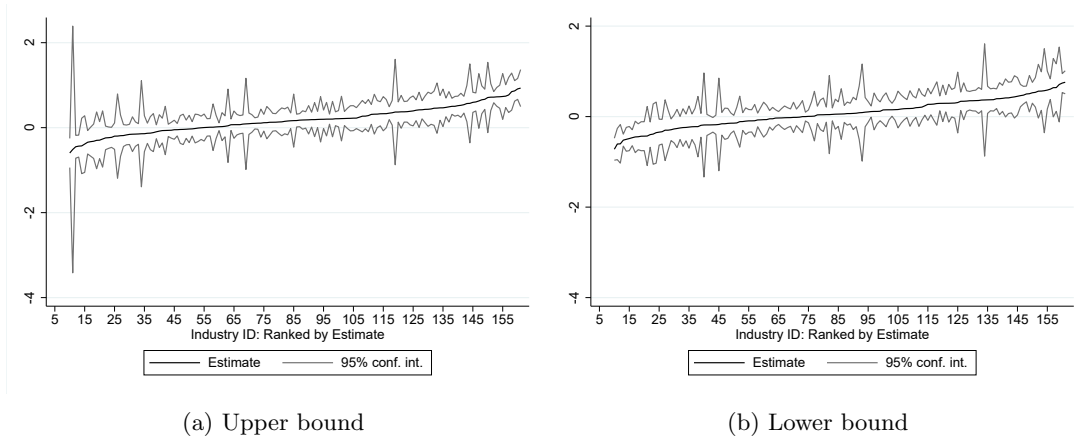


Figure B.1: Results by industry

**Note:** The figures show the distribution of point estimates (black lines) and 95% confidence intervals (grey lines), for industries in agriculture, mining, manufacturing, and services. Standard errors are clustered by country pair. The estimates are based on the ITPD-E database. The estimations on the left do not control for border-year dummies (“upper- bound”). The estimations on the right control for border-year dummies (“lower-bound”).