A sensitivity analysis on the impact of regional trade agreements in bilateral trade flows*

Un análisis de sensibilidad sobre el impacto de los acuerdos comerciales regionales en los flujos bilaterales de comercio

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Abstract

We estimate the effect of RTAs on bilateral exports by means of a gravity model analyzing its sensitivity to different specifications and methods. RTAs generate a sizable positive effect. However, shifting to country-pair and time-varying fixed effects systematically reduces coefficients. Nevertheless, the RTA effect is consistent across methods and specifications.
The RTA effect attributable to particular trade agreements displays high variability. While most RTAs increase trade, others present non-significant or negative results. We apply robustness checks to individual RTA estimates by presenting PPML time-invariant fixed effects and next to these, country-pair and time-varying fixed effects estimates. Thus, 38.2% of RTAs are positive and significant in both specifications. RTAs trade creation effects tend to prevail over trade diversion effects.

Key words: *International Trade, Trade Liberalization, Regional Trade Agreements RTA, Gravity Model, Economic Integration.*

JEL Classification: *F13, F14, F15, F53, F55.*

* The authors are grateful to the Editorial Board and two anonymous referees for their comments and corrections that greatly contributed to enhance the quality of this paper.
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Received: Novembre, 2019. Accepted: April, 2020.

Resumen

En este artículo estimamos, mediante un modelo de gravedad, el efecto de los Acuerdos Comerciales Regionales (ACR) en las exportaciones bilaterales, y realizamos un cuidadoso análisis de sensibilidad, considerando diferentes métodos y especificaciones. Los ACR presentan, por lo general, un efecto positivo considerable. Este impacto se reduce substancialmente al incluir efectos fijos país variables en el tiempo y efectos fijos individuales. No obstante, el efecto de los ACR es consistente a través de los métodos y especificaciones aplicados. Cuando el impacto de los ACR es calculado para cada acuerdo en particular, los coeficientes presentan una alta variabilidad. La mayoría de ACR presenta un impacto positivo. Otros presentan resultados no significativos o negativos. Para una mayor robustez de los resultados, los impactos de los ACR particulares fueron estimados con efectos fijos invariables en tiempo, y también con efectos fijos variables en el tiempo bajo el método de PPML. Así, el 38,2% de los ACR son positivos y significativos en ambas especificaciones. A su vez, los efectos de creación de comercio tienden a prevalecer sobre los efectos de desviación del comercio.

Palabras clave: *Comercio Internacional, liberalización económica, Acuerdos Comerciales Regionales ACR, modelo de gravedad, integración económica.*

Clasificación JEL: *F13, F14, F15, F53, F55.*

1. Introduction

After its creation in 1995, the World Trade Organization has been able to convince most of the countries to abide by the rules of multilateral trade. Nevertheless, its rounds of negotiations have come to a deadlock, partly explained by the difficulty of making agreements among too many countries of a heterogeneous nature. In the midst of this, Regional Trade Agreements RTAs, appeared as a more effective way to close trade deals. They presented exponential growth from the 80s to the first decade of the current century, to slow down in recent years. The question, then, arises about its effectiveness.

Trump's administration has dispensed with more than 70 years of liberal tradition in the United States, by dumping the Trans-Pacific Partnership (TPP) and the Transatlantic Trade and Investment Partnership (TTIP), the two most ambitions RTAs ever negotiated, while initiating a frontal trade war with China, and imposing duties on aluminium and steel worldwide. In the same direction, the United Kingdom has officially divorced from the most admired and profound RTA, the European Union. In a time where liberal ideas are under strain, answering the question of whether RTAs really increase trade is even more important. Despite substantial progress to compute RTA estimates, the debate about the effectiveness of RTAs remains open. The main objective of this paper is then to help answer the question: To what extent are RTAs able to create trade?

To do it, we employ the widely accepted approach of the gravity model. We build on the works of Baier and Bergstrand (2007), Martínez-Zarzoso *et al.* (2009); Kohl (2014) and Baier *et al*. (2019).

The main contribution of this paper is the presentation of a comprehensive sensitivity analysis based on a battery of relevant regression methods and specifications applied to the gravity equation on an updated database, providing the possibility of easy visualization and comparison, see Table 1. This information will also be valuable for future meta-analysis studies about the effect of RTAs on trade. We subsequently explore the RTA effect on bilateral trade for an ample sample of particular RTAs. Thus, coefficients for 123 particular RTAs comparing PPML time-invariant fixed effects (TIFE) and time-varying fixed effects (TVFE) estimates are presented in Table 3, which enables us to carry out robustness checks on their effectiveness. As far as we know, we are the first to present this comparison and analysis at disaggregated level for a large number of RTAs over a long period, see Table 4. Finally, we present results on trade creation and trade diversion for 25 relevant RTAs in Table 5.

Results from our most relevant specifications and methods point to a positive and significant effect of RTAs between 4.7% and 51.3% on bilateral exports. Considering particular RTAs, their impact is predominantly positive and significant. Trade creation effects in most of the cases offset trade diversion effects.

Gravity model estimations define what should be the normal pattern of trade, and then enable us to seek deviations from it, originated, for example, in the implementation of institutional arrangements. Given the counterfactual it offers, and its widespread use, the gravity model is tenable for calculating outcomes such as the expected gains from the entry into force of an RTA, or other institutional changes.

One important advantage of gravity models according to Bussière (2009) is that their results stem not only from a measure of multilateral trade integration (a country against all its trading partners), but also of bilateral trade integration (a country and each of its trading partners).

Our interest in finding the effects of RTAs in bilateral trade flows, hinges on the belief that higher international competition leads to greater productivity and higher cross-border exchanges increase wellbeing. We do not intend to disentangle this effect, although from Sachs *et al*. (1995), Wacziarg and Welch (2008) we have evidence that international trade promotes economic growth and then wellbeing. Similarly, Halpern *et al*. (2015) have established a positive relationship between firm import input access and productivity in the Hungarian economy, and Bas and Ledezma (2010) provided evidence of trade barriers reduction and with-in plant productivity increases in Chile.

In 1980, the GATT counted up to 83 signatories. In 2020 the number practically doubles, reaching 164 countries, now under the label of the WTO. Hayakawa and Kimura (2015) found that free trade agreements (FTAs) successfully reduce tariff rates and non-tariff barriers (NTBs). Nevertheless, mixed effects were reported by Afesorgbor (2017), Caporale *et al*. (2012), Didia *et al*. (2015), Kahouli & Maktouf (2015), Martin-Mayoral *et al*. (2016) who studied the impact of RTAs on exports by trade blocs in different regions as the Americas, Africa or Europe. Their results maintain alive a long-standing debate on the optimal mechanism for liberalizing international trade, confronting the multilateral negotiation approach to RTAs.

It is expected that membership to multilateral trade institutions would bear a strong positive effect on trade. Strong evidence for a positive WTO membership effect was found by Rose (2005), Subramanian and Wei (2007) and Kim (2011). Nevertheless, Eicher and Henn (2011) found evidence of an attenuated WTO membership impact after preferential trade agreements had entered into force. In view of the historical importance of this institution, this paper controls for country membership status in the WTO.

In parallel, the number of physical RTAs in force has steadily grown from 1980 to 2019. There were only 15 RTAs in 1980. They rose to 51 in 1995, 137 in 2005 to reach the number of 303 in 2019. Despite a slowdown in the number of new RTA negotiations worldwide, more RTAs are expected see the light in the years to come.

Special attention has been paid to Baier and Bergstrand (2007) who, using panel data on five years intervals, found that the average treatment effect of an RTA implies an increase of bilateral exports around 100% in 10 years. Another important contribution came from Magee (2008) who let the RTA dummy take leads and lags, thus finding significant anticipatory and slow motion impacts. Thus, in the long-run, an RTA increases trade on average by 89%. Regarding dynamics, Martínez *et al*., (2009), remark that bilateral exports are persistent and find significant effects for the lagged bilateral export flows, as well as for RTA coefficients at the disaggregated level.

RTA estimates have recently been reviewed downwards, a result that we confirm in this paper. This erosion effect was detected by De Sousa (2012) who focused on the effect of currency unions, and later by Kohl (2014) applying the Baier and Bergstrand's technique where he found that RTAs increased trade by at most 50%. Proving the reasons behind this behaviour goes beyond the scope of this paper. Yet, some hypothesis point out to the appearance of diminishing returns as more and more countries engage in RTAs; a rise in transaction costs coming from the multiplication of non-tariff measures such as rules of origin and local content requirements, or even a relaxation in the enforceability of existing RTAs due to a political movement of resistance to trade liberalization.

Following this introduction, the paper is organized as follows. Section 2 discusses methodological issues. Section 3 presents the data. Section 4 sets up the econometric specifications to be estimated. Section 5 presents and analyses results and section 6 concludes the paper.

2. Gravity model and methodology

Important advances in the micro-foundation of the gravity model are attributed to Anderson (1979) and Anderson and Van Wincoop (2003). They set up a model in which consumers maximize a homothetic Cobb-Douglas utility function that is identical in all countries; goods are differentiated by their country of origin, iceberg costs are assumed and only a fraction of the goods arrives at destination.

The mathematical approach developed by them puts multilateral resistance in the spotlight of the analysis. Their model takes us to estimate:

$$x_{ij} = \frac{y_i y_j}{y^w}\left(\frac{t_{ij}}{P_i P_j}\right)^{1-\sigma} \tag{1}$$

Where x_{ij} represents exports from country $_i$ to country $_j$; y^w, y_i and y_j represent world, country i and country j's GDPs, respectively; t_{ij} is a trade cost factor between i and j, consisting of geographical, political and institutional barriers. The parameter σ represents the elasticity of substitution between all goods and P_i and P_j are the multilateral resistance terms, which give us a measure of the relative openness of the economies. The Gravity equation is compatible with several underlying theories. A detailed discussion about the gravity model micro-foundation is available in Head and Mayer (2014).

Augmented gravity models control for confounders, which, if omitted, would bias the estimate of our parameter of interest on RTAs. Hence, we control for border contiguity and other cultural or institutional variables such as the use of a common language, Melitz and Toubal (2014), and colonial links, Head *et al*. (2010).

Anderson and van Wincoop (2003) pointed out the difficulties in estimating unbiased coefficients through cross-sections as well as the threat of omitted variable bias derived from multilateral resistance. Thus, Panel data models enabling for fixed effects specifications provided a solution to the fact that P_i and P_j, the so-called multilateral resistance terms in equation (1) are unobservable and the procedure to estimate them implies a non-linear routine. De Benedictis and Taglioni (2011) examine the sensitivity of OLS estimates to variations in fixed effects. These procedures control for endogeneity from unobservable heterogeneity and then for omitted variable bias derived from multilateral resistance. The authors consider the introduction of time-varying fixed effects for importing and exporting countries a robust solution.

Apart from the multilateral resistance difficulty, the possibility of endogeneity between bilateral trade and institutional trade liberalization variables is also prominent. Trefler (1993) pointed out that a country's decision to sign a regional trade agreement could not be completely exogenous. In the same way, Ghosh and Yamarik (2004) based on extreme bounds analysis showed that the RTAs coefficient computed with cross-sectional data could be biased in the presence of endogeneity and Baldwin and Jaimovich (2012) found that free trade agreements could be contagious (domino effect). When endogeneity is present, traditional estimation methods could result in inconsistent estimates. Instrumental variable methods can deal with endogeneity, allowing for stronger causal claims.

In that vein, Baier and Bergstrand (2007, 92) stated "standard cross-section techniques using instrumental variables and control functions do not provide stable estimates of RTA average treatment effect in the presence of endogeneity, and tests of over-identifying restrictions generally fail". They suggested that panel data methodologies must be implemented to estimate the RTA coefficient.

A panel approach will then be preferred over cross-section because it accounts better for country observed and unobserved time-varying or time-invariant heterogeneity. It provides the possibility of controlling for relevant relationships over time, avoiding the risk of choosing an unrepresentative year Antonucci and Manzocchi (2006). Panels also improve the efficiency of the estimates, Cheng Hsiao (2003). The panel structure would deal relatively well with the endogeneity problem considering that the reasons linked to RTAs not being exogenous should most probably be related to time-invariant heterogeneity (huge pre-existing trade flows, or contiguity).

Not all RTAs are equal. Dür *et al*. (2014) created deep integration indicators proving that differences on the depth of the agreements produces weaker effects for shallow agreements. Considering the heterogeneity of economic integration agreements, Egger and Nigai (2015) concluded that shifting to deeper trade agreement increases welfare, this effect being particularly high for some countries. Kohl *et al*. (2016); Ahcar and Siroën (2017) confirmed the effects of deep integration on trade, where deeper agreements result in larger gains. Baier *et al*. (2018) also found that certain integration settings produce greater impacts on the intensive margin than on the extensive margin.

Seeking better predictions of the effect of new economic integration agreements, Baier *et al*. (2018) and Baier *et al*. (2019) went beyond the importance of accounting for RTA heterogeneity. They found asymmetries in the RTA effect linked to the direction of trade. They also proved that country-pair heterogeneity is relevant as any given integration agreement can produce different effects on trade. For example, partners engaged in pre-existing economic integration agreements and distant pairs of countries obtain weaker gains out of further integration.

Considering that the main objective of this paper is to compare the average effect of an RTA through different specifications and methods, we want to make the caveat that not all RTA are equally designed. Hence, we do not expect to interpret these coefficients as precise predictions of the effect of any new RTA agreement, as literature acknowledges that information on RTA heterogeneity is required for accurate forecasting purposes. Nevertheless, to mitigate these shortcomings, we estimate the RTA effect for particular couples and blocks, where we can observe a long range of variability on the effect of RTA, possibly caused by this heterogeneity.

3. Data

To deal with the challenges mentioned above and to successfully estimate our variables of interest, this research set up an exhaustive data set to run a gravity model. It consists of bilateral trade flows for 153 countries from 1980 to 2018 that add up to 715.626 individual bilateral trade flows and an extensive set of control variables.

Bilateral Exports are taken in current dollars at fob values from the International Monetary Fund (IMF) Direction of Trade Statistics Database DOTS (2020). The current GDP in dollars, population in number of inhabitants and urban participation in percentages are provided by the World Development Indicators (WDI) database of the World Bank (2020). The surface in square meters as well as island and landlocked status were constructed by the author based on data from the World Factbook of the Central Intelligence Agency of the United States of America CIA (2020). Weighted distance in Km, common land border and colonial links stem from the CEPII (2013): Head *et al*. (2010) Gravity dataset.

The dummy variable for Regional Trade Agreements was constructed by the author based on the Regional Trade Agreements Information System (RTA-IS) of the World Trade Organization WTO (2020), and from de Sousa RTA data set for De Sousa (2012). Generalized System of Preferences GSP is built by the author based on the Database on Preferential Trade Arrangements of the World Trade Organization WTO (2020). The author based on the World Trade Organization WTO information (2020) constructed GATT membership and OECD membership based on information from the Organisation for Economic Co-operation and Development (OECD) (2020).

4. Econometric specifications

The equation to estimate with OLS, with time-fixed effects and exporter and importer time-invariant fixed effects is presented in (2) below:

$$lnX_{ijt} = \beta_0 + \beta_1 RTA_{ijt} + \psi_h \boldsymbol{S_{it}} + \phi_h \boldsymbol{M_{jt}} + \varphi_g \boldsymbol{Z_{ijt}} + \alpha_t + \alpha_i + \alpha_j + \varepsilon_{ijt} \qquad (2)$$

Where, the dependent variable lnX_{ijt} represents the natural logarithm of current dollar fob export values from country i to country j; β_1 is the RTA coefficient, our parameter of interest; β_0 is a constant term, α_t represents the time-fixed effects, α_i represents time-invariant exporter fixed effects, α_j are the importer time-invariant fixed effects and ε_{ijt} is an idiosyncratic error term.

Likewise, $\boldsymbol{S_{it}}$ and $\boldsymbol{M_{jt}}$ are vectors of time-varying monadic controls for exporters and importers respectively composed of h variables: $lnGDP_{it}$, $lnpop_{it}$, $urpart_{it}$, $OECD_{it}$ and $GATT_{it}$, $gspprovider_{it}$, $gspben_{it}$ as well as, $lnGDP_{jt}$, $lnpop_{jt}$, $urpar_{jt}$, $OECD_{jt}$ and $GATT_{jt}$.

Here, ψ and ϕ are vectors of coefficients to be estimated concerning the above control variables, and the subscript h indicates variables.

We define $lnGDP_{it}$ and $lnGDP_{jt}$ as the natural logarithms for current dollar GDPs from countries *i* and *j*; $lnpop_{it}$, $lnpop_{jt}$ are natural logarithms for the population in number of inhabitants of countries *i* and *j*; $urpart_{it}$ and $urpart_{jt}$ stand for the percentage of urban population in country *i* and *j* respectively; this could be seen as a measure of the degree of development of countries, as more developed countries tend to be relatively more urbanized.

Other non-dyadic variables attempt to control for institutional traits related to commerce; these are $gatt_{it}$ and $gatt_{jt}$ that take on 1 if countries *i/j* belong to the GATT/WTO respectively. We use variable $gspben_{it}$ that takes on 1 if country *i* is receiving the generalized system of preferences or any other unilateral preference scheme from country *j*, otherwise 0; $gspprovider_{it}$ takes on 1 if country *i* is granting the generalized system of preferences or any other unilateral preference scheme to country *j;* $oecd_{it}$ and $oecd_{jt}$ take on 1 if the countries *i/j* belong to the Organization of Economic Cooperation and Development OECD.

When no country fixed effects are introduced, controlling for time-invariant monadic variables such as the total surface of a country, the fact of being an island or being landlocked, helps to improve results. Then, vectors $\boldsymbol{S_{it}}$ and $\boldsymbol{M_{it}}$ are augmented with variables $lnarea_{it}$, isl_{it} and $landlocked_{it}$; and $lnarea_{jt}$, isl_{jt} and $landlocked_{jt}$ respectively. Here, $lnarea_{it}$ and $lnarea_{jt}$ are the natural logarithms for the surface in square km of country i and j; *Isl* takes on 1 if country i/j is an island, otherwise 0; and *landlocked* takes on 1 if country i/j is deprived of a direct access to the sea, otherwise 0.

Finally, $\boldsymbol{Z_{ijt}}$ is a vector of dyadic variables that helps to minimize possible bias, composed of g variables: $contg_{ijt}$, $comlang_{ijt}$, $col45_{ijt}$ and $lndist_{ijt}$ and φ is a vector of coefficients to be estimated concerning these dyadic variables; the subscript *g* is to indicate variables, where $lndist_{ijt}$ is the natural logarithm for the weighted distance between countries *i* and *j*; $contig_{ijt}$ takes on 1 if there is a common land frontier between *i* and *j*, otherwise 0; $comlang_{ijt}$ takes on 1 if at least 9% of the pair population share the same language, otherwise 0; $col45_{ijt}$ takes on 1 if both countries were under a colonial relationship before 1945, otherwise 0; and finally our variable of interest rta_{ijt} takes on 1 if both countries share a free trade agreement, otherwise 0.

The equation to be estimated with random effects or with country-pair fixed effects is presented in (3) below. Here we follow (4) assumption.

$$lnX_{ijt} = \beta_0 + \beta_1 RTA_{ijt} + \psi_h \boldsymbol{S_{it}} + \phi_h \boldsymbol{M_{it}} + \alpha_t + \alpha_{ij} + \mathcal{E}_{ijt} \tag{3}$$

$$Cov(EV_{ijtg}\ ,\ \alpha_{ij}) = 0\ , \qquad t = 1,2,\ldots,T\ ;\ ij = 1,2\ldots,N\ ;\ g = 1,2\ldots,k. \tag{4}$$

Where EV stands for explanatory variables, (*ij*) represents the entities, *t* represents years, and *g* is to enumerate the explanatory variables.

α_{ij} represents country-pair fixed effect. For the traditional fixed effect model (within transformation) (4) assumption is modified to allow for a differential intercept for each country pair *ij*, then, a correlation between at least some of the explanatory variables and the country-pair fixed effects is permitted. See (5). This method does not allow controlling for time-invariant exporter and importer fixed effects at the same time, as the pair-fixed effects are collinear with country fixed effects. Thus, all time-invariant variables are dropped by the within transformation, Greene (2011).

$$Cov(EV_{ijtg}\ ,\ \alpha_{ij}) \neq 0\ ,\ t = 1,2,\ldots,T\ ;\ ij = 1,2\ldots,N\ ;\ g = 1,2\ldots,k. \tag{5}$$

Increasing acceptance to estimate gravity models is acknowledged to the Poisson Pseudo Maximum Likelihood estimator. This technique has been defended by Santos Silva and Tenreyro (2006; 2011) and Fally (2015) as the more reliable method to estimate the gravity equation because it deals with heteroscedasticity problems better than traditional OLS methods. Furthermore, in their work of 2011, they presented further evidence that the PPML estimator generates consistent estimates, even in the presence of a large number of zero values in the data set, a recurrent difficulty in gravity models.

(6) presents the PPML specification when we introduce year fixed effects and exporter and importer time-invariant fixed effects:

$$X_{ijt} = exp(\beta_0 + \beta_1 RTA_{ijt} + \varphi_g \boldsymbol{Z_{ijt}} + \psi_h \boldsymbol{S_{it}} + \phi_h \boldsymbol{M_{jt}} + \alpha_t + \alpha_i + \alpha_j)\, u_{ijt} \tag{6}$$

Here, X_{ijt} represents the value of the fob merchandise exports from country i to country j in current dollars and $u_{ijt} = \exp((1-\sigma)\,\varepsilon_{ijt})$. We chose this specification to evaluate trade diversion for a set of interesting RTAs. Thus we introduce a vector of $\mathbf{RTA}_{it}$ trade diversion dummies next to their associated vector of $\mathbf{RTA}_{ijt}$. The subscript k stands for the number of RTA dummies included. (6) can now be read as:

$$X_{ijt} = exp(\beta_0 + \beta_k \boldsymbol{RTA_{ijt}} + \Upsilon_k \boldsymbol{RTA_{it}} + \varphi_g \boldsymbol{Z_{ijt}} + \psi_h \boldsymbol{S_{it}} + \phi_h \boldsymbol{M_{jt}} + \alpha_t + \alpha_i + \alpha_j)\, u_{ijt} \tag{7}$$

Below in (8) we relax the assumption of the maintenance of unchanging gaps among different intercepts, or stable tendencies, through time. The inclusion of time-varying country fixed effects in the PPML specification leads us to estimate.

$$X_{ijt} = \exp(\beta_0 + \beta_1 RTA_{ijt} + \varphi_g \boldsymbol{Z_{ijt}} + \alpha_{it} + \alpha_{jt}) u_{ijt} \tag{8}$$

Where α_{it} stands for time varying exporter fixed effects and α_{jt} are the importer time-varying fixed effects. In (9) we include country-pair fixed effects in a specification that not only control for time varying unobserved heterogeneity at the country level but also for time invariant unobserved heterogeneity at the individual level. This is the literature preferred specification.

$$X_{ijt} = \exp(\beta_0 + \beta_1 RTA_{ijt} + \varphi_g \boldsymbol{Z_{ijt}} + \alpha_{it} + \alpha_{jt} + \alpha_{ij}) u_{ijt} \tag{9}$$

Improvements in Stata procedures documented by Correia *et al*. (2019) and Larch *et al*. (2019) have made possible the estimation of models with larger number of fixed effects with the PPML estimator, such as those presented in equations (8) and (9).

5. Results

In accordance with Baldwin and Taglioni (2006) this paper includes specifications that control for the passing of time using time-fixed effects. This approach allows us to work properly with GDP dollars, avoiding the so-called bronze medal mistake, which occurs when deflating these time series to obtain their real values. Non-averaged bilateral trade data to avoid the silver medal mistake is also used. The inclusion of time-invariant country fixed effects permits the partial offsetting of the endogeneity problem caused by omitted variables, in what is known as the gold medal mistake.

Under the OLS and PPML method, this paper also controls for time-varying country fixed effects for importers and exporters. This procedure would furnish a robust estimate of RTAs that controls for multilateral resistant and other omitted variables that change with the passing of time. The summary of the results will be presented in Table 1 to make comparison easier.

5.1. Traditional methods of estimation

5.1.1. *Pooled ols specifications results*

In its first row, Table 1 presents results based on the pooled OLS specifications. An analysis of the RTA coefficients shows that the model with no fixed effects in column 1 estimates a rise of 39.0%, ($e^{0.329}$ -1) in bilateral exports affected by RTAs relative to flows not influenced by them. It underestimates the impact of RTAs on international bilateral trade with respect to other OLS models that control for fixed effects, excepting for model 8 which simultanusly control for TVFE and country-pair fixed effects. This specficiation indicates a rise of 28.5% in bilateral trade.

When only time-fixed effects are controlled for, model 2 on the pooled OLS specification, the RTA coefficient overreacts, see column 2 of Table 1, producing a rise of 104.6% in bilateral exports affected by a RTA with respect to bilateral trade not affected by RTAs. This is the highest global RTA estimate computed in this paper.

5.1.2. *Random effects and country-pair fixed effects results*

A random effects model assumes that unobserved individual effects are uncorrelated with the explanatory variables Wooldridge (2012). The random effects model moves RTA estimates downward with respect to pooled OLS, yet a positive and significant effect is persistent.

The RTA random effects estimates in model 6 (see column 6 and random effects row) imply a slightly less important increase in trade than the model 2, controlling for time-fixed effects but omitting time-invariant fixed effects. Thus, when time-fixed effects and exporter and importer time-invariant fixed effects are introduced together as in model 6, we obtain an increase of about 31.3% in bilateral exports. To distinguish which model performs better between OLS and random effects we applied Breusch and Pagan (1980) test that checks if random effects are present. Based on their Lagrangian multiplier test for random effects, the OLS pooled model is outperformed by the random effects model.

When we relax the assumption that country-pair individuals' effects are uncorrelated with covariates, we obtain a fixed effect model. This model creates fixed effects for each bilateral export flow that remains invariant through time. Thus, observed and unobserved time-invariant heterogeneity at the country-pair level is kept at bay.

The fixed effects model in column 3 of the Fixed Effects row estimates that bilateral exports sharing a RTA increase by 19.1%, relative to flows without RTAs. Introducing time fixed effects to this model results in an increase of 18.9% in bilateral trade.

Results from the Hausman's specification test, establish that the fixed effect model fits better than the random effects. Particularities at individual level are then correlated with the explanatory variables. Nevertheless, the fixed effect regression at the individual country-pair level generates estimates that could be underestimating the RTA effect on bilateral trade, particularly when time fixed effects are accounted for.

5.2. Current methods of estimation

5.2.1. *PPML specification results*

The PPLM seems to be the more reliable method to estimate the gravity model. Martínez-Zarzoso (2013) validates this method through a series of Monte Carlo experiments. Fally (2015), Montenegro *et al*. (2011) and Martín-Montaner *et al*. (2014) gives additional support to PPML estimations over another techniques.

In the specification without time-fixed effects or country fixed effects, Column 1 in Table 1, the PPML estimate of RTA presents a rise of 10.5%. The introduction of time-fixed effects and country time-invariant fixed effects corrects PPML estimates upwards.

Likewise, PPML estimations shift upward to the introduction of time-varying fixed effects for exporter and importer countries, see column 7. The coefficient is lower in the TIFE and time-fixed specification, column 6, than in column 7, by 0.074 points, equivalent to 10.8 percentage points, which is a relevant difference that deserves attention. The TIFE and time-fixed effect model estimated by PPML produces an increase of 40.5% in bilateral exports affected by a RTA, compared with bilateral export flows that do not profit from any RTA; the comparable result using TVFE estimated by PPML is 51.3%.

We observe a sharp reduction of the RTA estimate when using PPML controlling simultaneously for TVFE and country-pair fixed effects, see column 8. This behaviour is also present under the OLS method, but PPML accentuates the decline. Sharing an RTA will only increase trade by around 4.7% in this specification, which is the preferred approach in literature.

5.2.2. *The Baier and Bergstrand method*

The Baier and Bergstrand technique consists of controlling for multilateral resistance and RTA endogeneity by the means of introducing country-pair fixed effects and time varying fixed-effects on a panel of non-successive years that we call periods. In accordance with Baier and Bergstrand we estimated our model keeping 10 periods, so we retain information for intervals of four years. Baier and Bergstrand (2007) kept data for intervals of five years[1]. In model 8 of the Baier and Bergstrand specification where country-pair and time-varying country fixed effects are computed together, the introduction of a RTA will increase bilateral exports by around 30.0% for OLS and 3.9% for PPML. RTA estimates under the Baier and Bergstrand method, where country fixed effects are accounted for; present higher values compared with those including country-pair fixed effects specifications. See Table 1.

The Baier and Bergstrand method simplifies the analysis of the dynamics of RTA through time. Variable rta_{ijt-1} and rta_{ijt-2} will capture the impact of RTAs on bilateral exports four and eight years before their entry into force or phase-in effects. It also enables the evaluation of anticipatory effects. Thus, rta_{ijt+1} describes the effects of the announcement and pre-entry into force of RTAs. Table 2 presents results for OLS and PPML regressions based on the time varying fixed effects and country-pair fixed effects specification.

Using the Baier and Bergstrand method with OLS, the first lag of the RTAs is positive and significant, but this does not hold for PPML specifications. Introducing a second lag in the specification drops the significance of the first lag. Using OLS, the RTAs effect experience a cumulative increase of 30.5% in bilateral exports during the first four year of entry into force, and approximately a 32.3% cumulative effect during the eight first years. The four years prior to its entry into force, known as the anticipatory effect of RTAs, is non-significant, which suggest strict exogeneity of RTAs, mitigating doubts about reversal causality where an increase in trade could cause RTA appearance. Estimates using PPML show an absence of cumulative and anticipatory effects. See columns (5-8) in Table 2.

[1] The results we show are estimated with data for years 1980, 1984, 1988, 1992, 1996, 2000, 2004, 2008, 2012 and 2016.

5.3. Instrumental variable methods and dynamics

5.3.1. *The Hausman and Taylor instrumental variable estimator*

To deal with RTA endogeneity problem of the type suggested by Baldwin and Jaimovich (2012) where a new free trade agreement between A and B increases the probability that C will sign a RTA with A or B or due to pre-existing overtrading patterns raised by Baier and Bergstrand (2007), we resort to the Hausman and Taylor estimator. These authors have proposed an instrumental variables estimator that uses only the information within the model by taking deviations from group means that can be used as instrumental variables. (Greene, 2011). The correct use of instrumental variable methods requires data on a sufficient number of instruments that are both exogenous and relevant. Swamy *et al.* (2015) argue that such instruments, weak or strong, are often impossible to find.

The Hausman-Taylor estimator assumes that some of the explanatory variables are correlated with the individual-level random effect α_{ij}, while none of the explanatory variables are correlated with the idiosyncratic error ε_{ijt}. In addition to standard assumptions, we also assume that RTA was the only endogenous variable in the model. Through Monte Carlo simulations, this estimator has proved to be robust for endogenous time-varying variables in large sample and perfect knowledge gravity model frameworks. (Mitze, 2010).

H-T estimates of RTA indicate an increase in bilateral trade between 18.2% and 30.2%. We should take these estimates with prudence as the Hausman test applied indicate that we should prefer the fixed effects estimator over the Hausman and Taylor estimator. Nevertheless, being the Hausman and Taylor estimator an instrumental variable estimator we should gain some confidence on making causal claims about the RTA effect.

Other instrumental variable techniques were also considered. The instrumental variable fixed effects and random effects estimators were computed using as instruments for RTA its lags on *t*-3 and *t*-4, under the assumption that these variables only influence bilateral exports by the influence they exert on the variable of interest RTA. Results suggest an upward bias for RTA in the IV-Dynamics using the third and fourth lags of the RTA, compared to H-T estimates.

5.3.2. *GMM regression and the Arellano and Bond estimator*

The main purpose of the Arellano and Bond (1991) method is to consistently estimate the dependent variable lags. This technique also allows setting other explanatory variables as endogenous by using GMM-type instruments to compute the causal effects of endogenous covariates. We tried to take advantage of this possibility and intended to use it to correct a possible endogeneity bias on the RTA estimates, nevertheless the Sargan test and the Hansen test failed, suggesting that lagged RTA instruments and lagged bilateral exports instruments were not valid because of overidentification.

TABLE 1
RTA ESTIMATES SUMMARY

Fixed Effects Controls	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(9)
Time invariant Exporter Fixed Effects	NO	NO	NO	NO	YES	YES	NO	NO
Time invariant Importer Fixed Effects	NO	NO	NO	NO	YES	YES	NO	NO
Time Varying Exporter Fixed Effects	NO	NO	NO	NO	NO	NO	YES	YES
Time Varying Importer Fixed Effects	NO	NO	NO	NO	NO	NO	YES	YES
Country-pair Fixed Effects	NO	NO	YES	YES	NO	NO	NO	YES
Time Fixed Effects	NO	YES	NO	YES	NO	YES	NO	NO
Econometric Method								
Pooled OLS	0.329***	0.716***			0.591***	0.631***	0.679***	0.251***
Random Effects	0.164***	0.296***			0.210***	0.272***		
Fixed Effects (within)			0.175***	0.173***				
PPML	0.100***	0.257***	0.067***	0.080***	0.336***	0.340***	0.414***	0.046***
Baier and Bergstrand (OLS)	0.148***	0.362***	0.170***	0.225***	0.607***	0.644***	0.706***	0.262***
Baier and Bergstrand (PPML)	0.086**	0.251***	0.041*	0.066***	0.330***	0.339***	0.410***	0.038**
IV: Hausman and Taylor	0.167***	0.264**			0.183***	0.246***		
IV-Dynamics: RTA lags 3 and 4	0.268***	0.473***	0.304***	0.402***	0.356***	0.451***		
IV-Dynamics: Arellano-Bond (GMM)	-0.247**	0.493***						

Note: *** p<0.01, ** p<0.05, * p<0.1. Robust standard errors.
Source: Elaborated by the authors.

TABLE 2
DYNAMICS BAIER AND BERGSTRAND REGRESSION ON 153 COUNTRIES FOR 10 PERIODS FROM 1980 TO 2016. OLS AND PPML ESTIMATORS

	(1) *lnxijt*	(2) *lnxijt*	(3) *lnxijt*	(4) *lnxijt*	(5) *Xijt*	(6) *Xijt*	(7) *Xijt*	(8) *Xijt*
rtaijt	0.251***	0.061***	0.059***	0.079***	0.046**	0.023	0.020	0.022
rtaijt-1		0.205***	-0.019	0.009		0.019	–0.002	0.002
rtaijt-2			0.240***	0.236***			–0.021	0.021
rtaijt+1				–0.021				–0.009
Constant	15.323***	15.331***	15.340***	15.352***	4.837***	4.850***	4.865***	4.838***
Observations	126,441	115,169	103,085	84,271	103,583	91,237	79,087	63,654
R-squared	0.878	0.885	0.892	0.902	0.957	0.958	0.960	0.959

Note: *** p<0.01, ** p<0.05, * p<0.1. Robust standard errors. All estimations included country-pair and time varying exporter and time varying importer fixed effects. Columns 1-4 for OLS. Columns 5-8 for PPML.
Source: Elaborated by the authors.

Among the techniques and methods applied in this paper, Arellano Bond specifications are the only ones to produce a significant and negative effect for the RTA dummy, although it disappears after dummy year inclusion. We present the Arellano-Bond RTA results in Table 1, assuming the RTA dummy as an exogenous variable and using the dependent variable lags as instruments to correct for the endogeneity of the first lag of the bilateral exports.

The results reported in Table 1 for Arellano-Bond RTA coefficients come from a regression that uses the first lag of the dependent variable as a regressor to account for the dynamics of the model, and all available dependent variable lags as instruments for this. In column 1 we show the result for the equation without year fixed effects, and in column 2 we control with time dummies. For robustness, we verified that the RTA coefficient is systematically negative in the regressions with no year effects as we reduce the number of lagged instruments, and overidentification problems also persist.

These results could be interpreted as evidence of the static approach strength over the dynamic approach, and also as an invitation for further research on dynamics of the dependent variable and RTAs. Some interesting works on dynamics for international trade gravity models can be found in Caporale *et al.* (2012), Didia *et al.* (2015) and Kahouli & Maktouf, (2015).

5.4. RTA estimates summary

Table 1 summarizes RTA estimates results, taking into account the econometric method and the fixed effect mix introduced. The methods used include pooled regression, random effects, fixed effects (within), PPML, Baier and Bergstrand for OLS and for PPML, Hausman and Taylor, Instrumental variables and dynamics models. Thus, in static models, the RTA coefficient is always positive and significant. Depending on the method and specification employed, statics models coefficients can vary from an estimate 0.038 in the Baier and Bergstrand-PPML method with TVFE and country-pair fixed effects, to levels as high as 0.716 that comes from the OLS pooled regression using only time fixed effects.

5.5. RTAs effects at the disaggregated level

The effect of particular RTAs computed by means of dummy variables for each scheme such as EU, NAFTA or MERCOSUR has been reviewed in Magee (2008), Eicher and Henn (2011) and Kohl (2014), Baier and Bergstrand (2019) among others.

Most of the preceding studies on the effects of particular RTAs are estimated by OLS techniques. This paper offers 123 RTA estimates based on PPML over a database across 153 countries and observations from 1980 to 2018, see Table 3. We apply a robustness check to our individual RTAs estimates by presenting PPML time-invariant fixed effects and next to them time-varying fixed effects estimates. On the time-invariant country fixed effects specification, we control for RTA membership other than the RTA of interest, distance between countries

TABLE 3
PPML ESTIMATES FOR A GROUP OF 123 REGIONAL TRADE AGREEMENTS FROM A 153 COUNTRIES 1980-2018 DATA SET

Agreement	Year	(1) TIFE	(2) TVFE	Agreement	Year	(1) TIFE	(2) TVFE
		RTA coef.	RTA coef.			RTA coef.	RTA coef.
ASEAN free trade area	1992	0.165***	–0.211***	EC–Jordan	2002	–0.279***	–0.237***
ASEAN-Australia	2010	0.352***	0.028	EC–Lebanon	2003	0.366***	–0.249***
ASEAN-India	2010	0.242***	0.009	EC–Mexico	2000	–0.161***	–0.204***
ASEAN-Korea	2010	0.799***	0.156***	EC–Moldova	2015	0.322**	0.229***
ASEAN-New Zealand	2010	0.402***	0.348***	EC–Morocco	2000	0.763***	0.149***
Australia-Japan	2015	0.652***	0.218***	ECOWAS	1993	1.041***	0.697***
Australia-Korea	2015	0.758***	0.111	EC–Peru	2012	0.224***	–0.044
Australia-New Zealand	1983	1.269***	0.486***	EC–South Africa	2000	0.485***	–0.446***
Australia-Singapore	2003	0.274***	0.135**	EC–Syria	1977	0.515***	–1.139***
Australia-Thailand	2005	0.777*	0.483***	EC–Tunisia	1998	0.913***	–0.003
CAN (Andean Community)	1988	0.928***	0.956***	EC–Turkey	1996	0.462***	0.106***
Canada Colombia	2012	–0.524***	–0.104	Ecuador–EC	2017	–0.006	0.139**
Canada-EC	2018	–0.236**	0.163***	EFTA–Israel	1993	0.354***	–0.599***
Canada-EFTA	2009	0.226**	–0.261***	EFTA–Korea	2006	0.437***	–0.009
Canada-Jordan	2013	–0.805***	0.563***	EFTA–Peru	2012	1.574***	–0.016
Canada-Peru	2009	0.857***	0.487***	GCC	2003	–0.812***	0.032
CEFTA	2007	0.389***	0.181***	Group of Three	1995	0.504***	0.895***
Chile Colombia1	1994	0.905***	–0.025	India–Japan	2011	–0.685***	–0.140***
Chile Colombia2	2009	0.924***	0.106	India–Malaysia	2011	0.436***	–0.029
Chile-Australia	2009	–0.852***	0.524***	India–Singapore	2005	0.300***	–0.053
Chile-China	2006	1.494***	0.606***	India–Sri Lanka	2001	1.251***	0.532***
Chile-EC	2003	0.201***	–0.247***	Japan–ASEAN	2008	0.581***	0.108***
Chile-India	2008	0.669***	–0.021	Japan–Indonesia	2008	0.648***	–0.111***
Chile-Japan	2007	0.867***	0.274***	Japan–Malaysia	2006	0.661***	0.121***
Chile-Korea	2004	1.579***	0.246***	Japan–Mexico	2005	–0.099	0.034
Chile-Malaysia	2012	–0.919***	0.035	Japan–Mongolia	2016	–0.355	0.736***
Chile-Peru1	1999	0.819***	0.250***	Japan–Peru	2012	0.318**	0.056
Chile-Peru2	2009	0.657***	–0.292***	Japan–Philippines	2008	0.507***	0.294***
Chile-Thailand	2015	0.120	0.333***	Japan–Singapore	2002	0.250***	0.084**
Chile-Turkey	2011	–0.297***	0.209**	Japan–Switzerland	2009	0.654***	0.100
Chile-Vietnam	2014	0.719***	0.431***	Japan–Thailand	2007	0.885***	0.237***
China-ASEAN	2005	–0.160***	0.146***	Japan–Vietnam	2009	0.636***	–0.050
China-Costa Rica	2012	0.075	0.229	Korea Republic–Canada	2015	–0.033	–0.039
China-New Zealand	2008	0.188*	0.506***	Korea Republic–India	2010	0.058	–0.023
China-Pakistan	2007	–0.096	0.181***	Korea Republic–New Zealand	2016	0.156***	0.030
China-Peru	2010	1.351***	0.339***	Korea Rep.–Singapore	2006	0.649***	0.300***
China-Singapore	2009	–0.256***	–0.025	Korea Republic–Turkey	2013	0.424***	0.301***
CIS	1994	1.561***	0.020	Korea–Peru	2012	1.292***	0.348***
COL (CAN) MERCOSUR	2005	–0.006	0.247***	Mauricio–Turkey	2013	0.865***	0.833***
Colombia Northern Triangle	2009	0.514***	0.489***	MERCOSUR	1991	1.139***	0.592***
Colombia-Costa Rica	2017	–0.257	–0.220*	MERCOSUR–India	2009	0.325***	0.206**
Colombia-EC	2013	0.185**	–0.0317	Mercosur–Peru	2006	0.024	–0.105**
Colombia-EFTA	2011	0.345**	–0.396***	NAFTA	1994	0.871***	0.439***
Colombia-Korea	2017	0.369***	0.036	PAFTA	1998	–0.699***	0.547***
COMESA	1994	1.273***	1.067***	SAFTA	2006	0.284**	–0.041
EAEC	1997	1.155***	0.417***	Southern African Develop. Comm.	2000	1.992***	0.229***
EC Enlargement (10)	1981	0.239***	0.018	Turkey–EFTA	1992	0.031	–0.268***
EC Enlargement (12)	1986	0.259***	0.106***	Ukraine–Belarus	2006	1.658***	0.519***
EC Enlargement (15)	1995	0.309***	0.083***	Ukraine–Kazakhstan	1998	1.864***	0.188
EC Enlargement (25)	2004	0.277***	0.086***	Ukraine–Turkmenistan	1995	3.266***	0.366
EC Enlargement (27)	2007	0.496***	0.143***	US–Australia	2005	–0.781***	–0.237***
EC Enlargement (28)	2013	0.534***	0.076***	US–Bahrain	2006	–0.163	0.089
EC-Albania	2006	0.976***	–0.031	US–CAFTA–DR	2006	0.477***	0.132***
EC-Algeria	2005	0.262***	–0.227***	US–Chile	2004	–0.228**	0.324***
EC-Cameroon	2009	0.433***	–0.464***	US–Colombia	2012	0.153**	0.026
EC-Caricom	2008	–0.623***	–0.122***	US–Israel	1985	1.093***	0.253***
EC-Côte d'Ivoire	2009	0.275***	–0.283***	US–Jordan	2001	0.339***	0.778***
EC-Croatia	2002	0.803***	–0.130***	US–Morocco	2006	–0.716***	0.367***
EC-EFTA	1973	0.220***	–0.082***	US–Oman	2009	–0.703***	0.302***
EC-Egypt	2004	0.185***	–0.342***	US–Peru	2009	–0.147	0.096**
EC-Ghana	2017	–0.416	–0.472***	US–Singapore	2004	–0.021	–0.273***
EC-Israel	2000	0.042	–0.283***				

Source: Elaborated by the author. *** p<0.01, ** p<0.05, * p<0.1. Robust standard errors.

Note: Columns (1) are estimated with time-invariant fixed effects and time fixed effects. Columns (2) include time-varying fixed effects and country-pair fixed effects.

i and *j*, common land frontier between *i* and *j*, if the country-pair shares the same language, and if both countries were under a colonial relationship before 1945. Profiting from recent PPML computing power improvements, Correia *et al* (2019), we respectively estimate PPML country-pair and time-varying fixed effects for individual RTAs. This is one of the major contributions of this paper.

As can be seen in Table 4, column 1, most of the RTA estimates, 94 out of the 123, equivalent to 76.4% of the sample, show a positive sign. Column 2, gives a less optimistic view, presenting 80 positive estimates. In addition, 62 RTA estimates bear out positive signs in both specifications, and 47, equivalent to 38.2% are positive and significant in both PPML specifications. These results, point to a larger proportion of trade agreements that are successful in promoting trade than in Kohl (2014), who reported that only 44 out of 166 RTAs, equivalent to 26.5% of their sample presented a positive and significant effect. Another interesting comparison is Baier *et al*. (2019) who used a sample of 65 RTAs and found positive statically significant effects for the majority of the agreements, 54%. Nevertheless, their results are not strictly comparable to ours, as they account for the effect of lagged RTAs.

The median RTA on this sample increases trade by 42.2%, ($e^{0.352} - 1$). Despite the dispersion, around 75.6% of RTA's estimates fall within one standard deviation of this median effect and 93.5% within two standard deviations.

Some straightforward outliers are the Chile-Malaysia, Gulf Council Countries GCC, PAFTA and EC-Caricom agreements, which seem to be highly counterproductive to trade creation, while the largest positive effects are posted by Ukraine-Turkmenistan, SADC, the Chile-Korea, EFTA-Peru and the Ukraine-Kazakhstan agreement. The latest impressive results of these cases concern former Soviet Union countries and could be attributed to some kind of transition effect or measurement error that could bias their estimates upward. Chile-Malaysia and GCC, Ukraine-Turkmenistan, EFTA-Peru and Ukraine-Kazakhstan become non-significant under the TVFE and country-pair specification. As we can observe in Table 4, around 23% of the RTA lose significance when shifting from TIFE to TVFE. The number of agreements significant in both specifications is 77, equivalent to 63%. One intriguing result is that only 50.4% of RTAs are positive and significant under the country-pair and TVFE specification. That number is substantially higher under the TIFE specification reaching 71.5% of RTAs.

Considering results in both specifications, United States agreements present mixed results, showing trade creation with Israel, Jordan and Colombia while the agreement with Bahrein is non-significant. Counterproductive effects appear with Australia. Similarly, European Union agreements outside its zone tend to produce mixed results. Particularly successful seem to be the agreements with Albania, Turkey and Moldova. Agreements with CARICOM, Jordan and Mexico present significant negative effects in both specifications. On the other side of the Pacific Ocean, 64% of the RTAs signed by Japan show a positive sign, and only its agreement with India produces a negative impact. China's RTAs tend to promote trade. Robust results are present in its agreements with Chile and Peru.

TABLE 4
DESCRIPTIVE STATISTICS FOR 123 RTA COEFFICIENTS ESTIMATED BY PPML ON A GRAVITY MODEL

	Specifications	
	TIFE	TVFE
Total number of RTAs	123	123
RTA coefficients (average)	0.407	0.119
Number of RTAs presenting a positive effect	94	80
% of RTAs presenting a positive effect	76.4%	65.0%
Average of positive RTA coefficients	0.653	0.29
Number of RTAs presenting a negative effect	29	43
% of RTAs presenting a negative effect	23.6%	35.0%
Average of negative RTA coefficients	0.391	0.2
Number of (+ or -) significant RTA coefficients (below the 0.10 level of significance)	106	89
% of (+ or -) significant RTA coefficients	86.2%	72.4%
Average coefficient for significant RTAs	0.484	0.155
Number of positive (+) and significant RTA coefficients	88	62
% of positive and significant RTA coefficients	71.5%	50.4%
Average coefficient for positive and significant RTAs	0.694	0.352
Number of negative (-) and significant RTA coefficients	18	27
% of negative and significant RTA coefficients	14.6%	21.9%
Average coefficient for negative and significant RTAs	–0.541	–0.298
	Total	%
Number of RTAs losing significance by shifting from TIFE to TVFE	29	23.5%
Number of non-significant RTAs gaining significance by shifting from TIFE to TVFE	12	9.8%
Number of significant (+ or -) RTAs on both specifications (TIFE and TVFE)	77	62.6%
Number of non-significant (+ or -) RTAs on both specifications (TIFE and TVFE)	5	4.1%
Number of positive (+) and significant RTAs on both specifications (TIFE and TVFE)	47	38.2%
Number of negative (-) and significant RTAs on both specifications (TIFE and TVFE)	5	4.1%

Source: Elaborated by the authors. Note: (TIFE) stands for time-invariant fixed effects. (TVFE) stands for time-varying fixed effects.

A final caveat: RTA coefficients at the disaggregated level should be read with caution. The scope and depth of the agreements change considerably from one RTA to the other. In theory it could be expected that deeper agreements produce higher increases in cross-border flows than those which are shallow. Equally important is the enforceability of these arrangements, especially in the case of politically unstable developing countries.

5.6. RTAs trade creation or trade diversion

Following Ghosh and Yamarik, (2004) and Eicher, Henn and Papageorgiou (2012) we use two sets of dummy variables to pick up RTA trade creation and trade diversion effects. Trade diversion occurs if a trade block creates trade in detriment of more productive third countries excluded from the agreement. The first, RTA_{ijt}, in (7) implies that both trading partners are members of the same RTA, the second, $DivRTA_{it}$ indicates that one country, whereas exporter or importer is a member of the RTA we are estimating.

Ghosh and Yamarik, (2004) define $DivRTA_{it}$ as a vector of variables which measures current membership of either country *i* or *j* in a RTA and thus, captures the external effects of the RTA on trade with countries outside the zone. The coefficient Υ_k for $DivRTA_{it}$ is interpreted as a measure of lower or higher than normal trade between nations in the trading bloc, and a country outside the bloc relative to a random pair of countries.

Hence, a negative sign for Υ_k indicates less trade with non-members and is interpreted as evidence of trade diversion.

In this section, we select a group of 25 interesting RTAs to evaluate whether trade diversion is actually mitigating the impact of RTAs on trade. As in Magee (2008) our estimates point to trade creation effects for ASEAN, MERCOSUR and NAFTA. The following analysis will be based on results for the TIFE specification because some trade diversion effects could not be estimated under the TVFE specification, due possibly to collinearity problems as too many fixed effects were dropped to perform estimations.

Thus, a third of the agreements trade creation effects are mitigated by trade diversion effects. In 6 cases the intra-block trade creation effect is sufficiently strong to resist trade diversion as in Australia-Korea, Colombia-Northern Triangle, the Group of 3, ECOWAS, EC-Turkey and NAFTA. For half of the sample of analysed RTAs, the extra-block effect reinforces the intra-block trade creation effects. Conversely, the trade diversion effect outstrips the trade creation intra-block effect in ASEAN-Japan and adds to intra-block negative effects in Canada-Colombia, Chile-EC, EC-Israel and Peru-United States. See Table 5.

6. Conclusion

This paper examines the effect of regional trade agreements RTAs on international bilateral trade flows. Based on the gravity model, we perform a sensitivity analysis to the effect of the RTA dummy, applying a wide range of econometric methods and model specifications. Our database consists of an unbalanced panel for 153 countries, including observations from 1980 to 2018. Particular attention is given to Poisson Pseudo Maximum Likelihood

TABLE 5

TRADE CREATION AND TRADE DIVERSION: PPML ESTIMATES FOR A GROUP RTAS FROM A 153 COUNTRIES 1980-2018

Agreement	1 Intra-block effect *RTAijt*	2 Extra-block effect *DivRTAit*	3 Net effect	4 Intra-block effect *RTAijt*	5 Extra-block effect *DivRTAit*	6 Net effect
Australia - Korea Republic	0.436***	–0.029	0.407	0.114	0.028	0.14
Andean Community	0.809***	0.019	0,82	0.709***	.	
Association of Southeast Asian Nations ASEAN	0.200***	0.137***	0,36	–0.192***	.	
ASEAN - Japan	0.091**	–0.346***	–0,26	0.262***	0.105***	0.37
Canada - Colombia	–0.463***	–0.154***	–0,60	–0.122	.	
Canada - European Union	0.591***	0.092**	0.683	0.277***	0.038	0.32
Caribbean Community CARICOM	3.342***	0.108	3,45	0.828*	.	
Chile - Colombia	0.902***	0.203***	1,17	–0.195*	.	
Chile - European Union	–0.319***	–0.158***	–0,44	–0.373***	–0.051***	–0.42
Chile - Vietnam	0.534***	0.179***	0.355	0.031	–0.245***	–0.21
Common Market for Eastern and Southern Africa	1.234***	–0.019	1,21	–0.245	–0.671***	–0.92
Colombia - EFTA	0.153	0.097***	0,24	0.084	0.275***	0.36
Colombia-(Guatemala-El Salvador-Honduras)	0.433***	–0.008	0.813	0.437***	0.022	0.46
Colombia-Mexico-Venezuela (The Group of 3)	0.655*	–0.013	0,41	0.560***	.	
Colombia - United Estates	0.294***	0.05	0,32	–0.031	.	
Economic Community of West African States	0.963***	–0.093**	0,88	0.662***	.	
European Free Trade Association EFTA	0.665***	0.196***	0,86	–0.009	.	
European Union EU(27)	0.591***	0.084***	0,67	0.414***	0.055	0.47
European Union - Israel	–0.127	–0.055	–0,20	–1.288***	–0.511***	–1.80
European Union - South Africa	0.786***	0.140***	0,93	0.611**	0.516***	1.13
European Union - Tunisia	0.939***	0.012	0,94	0.051	0.026	0.08
European Union - Turkey	0.490***	–0.002	0,51	0.166***	0.002	0.17
Southern Common Market MERCOSUR	1.311***	0.044	1,36	0.556***	.	
North American Free Trade Agreement NAFTA	0.463***	–0.220***	0,25	0.405***	.	
Peru - United States	–0.328***	–0.106***	–0,43	0.068	.	
Observations	674133	674133		457250	457250	
R-squared	0.899	0.899		0.958	0.958	

*** p<0.01, ** p<0.05, * p<0.1. Robust standard errors. Source: Elaborated by the author.

Note: Columns (1) and (2) include time-invariant country fixed effects and time fixed effects. Columns (4) and (5) stand for time-varying fixed effects and country-pair fixed effects

(PPML), which is the method that, when applied with fixed effects, best seems to contend with heteroscedasticity problems and bias from a high proportion of trade flows registered as zero.

A strong positive impact for RTA is consistently found on most specifications. Once multilateral resistance and other unobserved variable bias are controlled by the introduction of time-varying country fixed effects in a PPML regression, we find that RTAs increase bilateral trade flows by 51.3%, with respect to those trade flows with no agreements. When country-pair fixed effects are added to the previous specification, the RTA effect is reduced to 4.7%, still economically significant, as it confirms that efforts to close international trade deals are fruitful.

RTA cumulative effects are found using OLS, but its significance disappears with the PPML method. Instrumental variable methods were also tested. Using the third and the fourth lags of RTAs as instruments for RTA, as well as, employing the Hausman and Taylor estimator that introduces instrumental variables to deal with endogeneity, gives sizable and significant results.

We found considerable variations in the estimates of RTAs at the disaggregated level. While most of these successfully increase trade, others seem to destroy it, or are non-significant. When only time-invariant fixed effects and time fixed effects were included, 71.5% of RTA were positive and significant; this number slid to 50.4% for the time-varying and country-pair fixed effects specification. Robustness checks for individual RTAs based on the comparison of the PPML time-invariant fixed effects specification and the time-varying and country-pair fixed effects specification show that 38.2% of RTAs are positive and significant in both specifications.

The wide range of individual RTA estimates [-1.139; 3.266] could be explained by the fact that RTAs are heterogeneous in scope and depth. Another hypothesis points to lack of enforceability, meaning that a number of RTAs are not completely implemented in practice and remain only as a written statement, a line of research worth exploring.

Trade diversion effects were computed for a sample of RTAs. At large, trade creation effects tend to be stronger than trade diversion effects or even be reinforced by an open trade block expansion effect. Nevertheless, the potentiality of RTA to improve well-being must not be given for granted, as in certain cases trade diversion is found to outstrip trade creation effects.

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APPENDIX

List of countries included in the gravity model database

Albania, Algeria, Angola, Argentina, Australia, Austria, Azerbaijan, Bahrain, Bangladesh, Barbados, Belarus, Belgium, Belize, Benin, Bermuda, Bolivia, Brazil, Brunei, Bulgaria, Burkina Faso, Burundi, Cambodia, Cameroon, Canada, Cape Verde, Central, African Republic, Chad, Chile, China, Colombia, Congo Democratic, Congo Republic, Costa Rica, Ivory Coast, Croatia, Cuba, Cyprus, Czech Republic, Denmark, Djibouti, Dominican Republic, Ecuador, Egypt, El Salvador, Equatorial, Guinea, Estonia, Ethiopia, Fiji, Finland, France, Gabon, Gambia, Georgia, Germany, Ghana, Greece, Grenada, Guatemala, Guinea, Guinea-Bissau, Guyana, Haiti, Honduras, Hong Kong, Hungary, Iceland, India, Indonesia, Iran, Iraq, Ireland, Israel, Italy, Jamaica, Japan, Jordan, Kazakhstan, Kenya, South Korea, Kuwait, Kyrgyzstan, Latvia, Lebanon, Liberia, Libya, Lithuania, Luxembourg, Madagascar, Malawi, Malaysia, Mali, Malta, Mauritania, Mauritius, Mexico, Moldova, Mongolia, Morocco, Mozambique, Nepal, Netherlands, New Zealand, Nicaragua, Niger, Nigeria, Norway, Oman, Pakistan, Panama, Papua New Guinea, Paraguay, Peru, Philippines, Poland, Portugal, Qatar, Romania, Russia, Rwanda, Samoa, Saudi Arabia, Senegal, Sierra Leone, Singapore, Slovakia, Slovenia, South Africa, Spain, Sri Lanka, Sweden, Switzerland, Syria, Tajikistan, Tanzania, Thailand, Togo, Tonga, Trinidad and Tobago, Tunisia, Turkey, Turkmenistan, Uganda, Ukraine, United Arab Emirates, United Kingdom, United States, Uruguay, Uzbekistan, Venezuela, Vietnam, Yemen, Zambia, Zimbabwe.