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Do Regional Trade Agreements Really Boost Trade? Estimates for Agricultural Products

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Highlights

- We document precisely liberalization patterns for 74 regional trade agreements (RTAs), over the period 1998-2009, for the agricultural and food sector.

- The mean elasticity of substitution across imports at the product level is estimated to be slightly below 4.

- RTAs have increased bilateral agricultural and food exports between partners by 30% to 40% on average.

- RTAs increase the probability to export a given a product to a partner country, but this impact is estimated to be lesser than one percentage point on average

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▍ Abstract

The trade effects of tariff preferences are assessed using difference-in-differences panel estimations, whereby exports to third destinations and imports from third origins are used as benchmarks. The method is applied at a detailed product level for 74 agreements, over the period 1998-2009, for the agricultural and food sector. We estimate the mean elasticity of substitution across imports at the product level to be slightly below 4, with significant but limited differences across types of agreements and level of preferential margin. Counterfactual simulations suggest that RTAs have increased bilateral agricultural and food exports between partners by 30% to 40% on average, with a marked heterogeneity across agreements. RTAs are also found to increase the probability to export a given a product to a partner country, but this impact is estimated to be lesser than one percentage point on average.

▍ Keywords

Regional trade agreement, international trade, agricultural products,tariff protection.

▍ JEL

F13, Q17.

Do Regional Trade Agreements Really Boost Trade?
Estimates for Agricultural Products

Sébastien Jean[*] & Jean-Christophe Bureau[**]

Introduction

Multilateralism has stalled since the implementation of the Uruguay Round ended in 2004, but regional trade agreements (RTAs) have proliferated: the World Trade Organization has received 612 notifications of RTAs as of 7 April 2015 and estimates that 406 were in force, up from 180 in 2005.[1] In the meantime, the share of world trade occurring within RTAs has been growing steadily, overreaching 30 percent even when intra-EU trade is excluded (Bureau et al. 2015).

While RTAs are likely to profoundly transform the established order of international trade, their actual impact on trade flows remains controversial. Literature surveys show that estimates are highly variable, demonstrating a lack of robustness (Head and Mayer 2014; Cipollina and Salvatici 2010; Ghosh and Yamarik 2004). This raises empirical questions about the actual consequences of a shift toward regionalism for global trade. Some authors find that RTAs generate large trade flows between signatories, albeit often at the expense of third countries (e.g., Grant and Lambert 2008; Egger and Larch 2011; Egger and Wamser 2013; Fugazza and Nicita 2013; Caliendo and Parro 2015). Others find that the impact of RTAs on trade flows tend to be lower than often expected, in particular because of the large number of goods subject to low duties (Carpenter and Lendle 2011) or because of the contrasting scope and depth of trade agreements (WTO 2011). That is, in spite of the growing literature, there is still need for robust estimates of the shift to regionalism and potential impacts of an RTA for a country that is considering concluding such an agreement.

To shed light on this question, this article takes advantage of the longitudinal variability in tariffs resulting from differences across products and across agreements, as well as from the progressive implementation of RTAs. Such approach requires very detailed approach, with tariff and trade measured in a consistent way for each product and each year, on a bilateral basis. To do this, we built a dataset covering 74 RTAs over the 1998-2009 period, which describes the progressive implementation of tariff cuts product by product (at the 6-digit level

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[1] In this article, the term RTAs is to be understood in a broad way, i.e. including preferential agreements between countries in different regions. Note that the WTO figures tend to overstate the number of active RTAs since those that cover goods and services liberalization are counted twice by the WTO (see Grant, 2013). Avoiding such double-counting, the WTO registered 262 "physical" agreements in force in April 2015.

of the Harmonized System) and year by year. Due to limitations in some of the sources used, this uniquely detailed dataset covers only agricultural products.

The benefit of this approach is to be able to assess the trade impact of tariff cuts only out of longitudinal changes, thus making it possible to control any time invariant factor, even when it is specific to exchanges of a given product, between a given pair of countries. As such, this estimation framework is fully consistent with a large array of theoretical frameworks (see Head and Mayer 2014 for developments on the required conditions); it accounts for multilateral resistance factors (Anderson and van Wincoop 2003), and endogeneity concerns are minimal, since any time invariant determinant is controlled for.

Consistent estimates based on the standard specification would require taking into account a number of fixed effects (exporter-importer-product, export-product-year, importer-product-year) which is intractable for such a detailed analysis covering a significant number of agreement:[2] with 35 countries, 700 products and 12 years covered, the corresponding number of fixed effects would be 35x35x700=857,700 for exporter-importer-product and 35*700*12=294,000 for exporter-product-year and for importer-product-year. We use instead a method that relies on reference groups to control for importer- and exporter-specific determinants, and we use a ratio-of-ratios approach at the product level. We extend a method developed by Romalis (2007) in two ways. First, we apply it to a large number of agreements rather than a bilateral comparison. Secondly, our estimation is based on a count model (Poisson), avoiding the bias inherent in the log-transformation of a heavily heteroskedastic model. In addition to Romalis' approach, we distinguish between the impact of these agreements on pre-existing trade flows and on the creation of new flows.

We find that trade agreements affect pre-existing trade flows with a mean elasticity of substitution at the product level of 3.7, so that a 1% preferential margin with respect to the MFN duty rate increases trade by slightly less than 4% on average, once the RTA is fully implemented. According to our counterfactual simulations, the RTAs studied would have increased bilateral agricultural and food exports by 30 to 40% on average. Differences across agreements are large though, with an almost zero impact in some cases, and a doubling of initial trade flows in others. Although statistically significant, the impact on the extensive margin, that is new trade flows, is very weak: a 1% preferential margin increases the probability of exporting by 0.05 percentage point. These results show how easy it is to underestimate the importance of RTAs on trade flows on the basis of the relatively small preferential margins in the agricultural and food sector. RTAs also include non-tariff provisions, but our estimates do not hint at a significant impact on bilateral trade flows on average, for products where tariffs are not cut.

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[2] The estimates presented below cover 35 countries, 700 products and 12 years. The corresponding number of fixed effects would thus be 35x35x700=857,700 for exporter-importer-product and 35*700*12=294,000 for exporter-product-year and for importer-product-year. In the present context of multiway, unbalanced panel, relying on within estimators would be inconsistent, meaning that these fixed effects would have to be taken into account explicitly.

1. Gravity models and their application to trade agreements

Gravity models are the most common approach used to measure the impacts of a trade agreement. General assumptions are sufficient to yield gravity-type equations, making them a flexible analytical framework. An extensive body of literature demonstrates the great interest from a practical point of view as well as the theoretical consistency of this approach (e.g., Baldwin and Taglioni 2006 and Head and Mayer 2014, among many others). The basic form of the gravity model is $X_{ijt} = G_t S_{it} M_{jt} \phi_{ijt}$, where the indices *i*, *j* and *t* denote the exporting country, the importing country, and the year, respectively; *X* denotes the value of the trade flow; G is a time-specific constant; S is a variable linked to the exporter's attributes and M to the importer's attributes; the influence of determinants of trade specific to each country pair is represented by ϕ. Once the relevant variables have been identified, the gravity model can be directly estimated, most often in its log form adding an error term.

In general, evaluations of the impact of RTAs on trade assume that the influence of the agreements can be separated out from other pair-wise specific trade determinants, and a dummy variable is introduced to capture the effect of RTAs. This analytical framework has given rise to a prolific literature. However, recent contributions to the literature have revealed that most of the older estimates were beset by errors. While some are easy to correct—such as the poor choice of a deflator or the flawed calculation of the mean unidirectional trade flows (Baldwin and Taglioni 2006)—others like multilateral trade resistance factors are more fundamental (Anderson and Wincoop 2003). So is endogeneity, the creation of RTAs motivated by missing variables, which also contribute to determining the intensity of trade (Anderson 2011; Baier and Bergstrand 2007, 2009). The standard use of dichotomic variables to indicate whether an agreement existed between the members also relies on stringent underlying assumptions. Grant and Lambert (2008) have shown that some of them were not realistic.

To avoid these estimation problems, the strategy adopted here is to transform the dependent variable, as proposed by Romalis (2007) in the case of a bilateral agreement. This transformation on the multiplicative structure of the model makes the estimate open to interpretation as difference-in-differences in their logarithmic form, as are the methods used in Hallak (2006) or Head et al. (2010), referred to in the latter as the *Tetrads* method. Application of this method requires both an exporter control group and an importer control group. By choosing as control groups countries whose trade policy toward signatories did not change during the study period, bilateral trade determinants with control groups can be assumed unchanged, so that movements in the difference in differences in trade flows can be explained by trade liberalization between the signatories.

Unit-by-unit division of the standard models for two exporting countries *i* and *i'* and for a given product *k*, yields: $R_{ii'jkt} = X_{ijkt}/X_{i'jkt} = (S_{ikt} \phi_{ijkt})/(S_{i'kt} \phi_{i'jkt})$, where *R* is the ratio of country *j*'s imports of *k* from suppliers *i* and *i'* respectively. This formulation allows the general term and the importer-specific term to be eliminated. If we let exporter-specific attributes be invariant (or vary at the same relative rate regardless of the sector), this

equation allows the evolution over time of the determinants of the bilateral intensity of trade ϕ to be identified, provided the bilateral determinants of trade are constant in the case of partner i'. Exporters attributes that change over time can also be controlled by examining the relative volume of imports from suppliers i and i' (labeled *ROR*, for Ratio of Ratios) on markets j and j', obtained by dividing equation (1) unit-by-unit for each of market j and j':

(1) $ROR_{ii'jj'kt} = \dfrac{R_{ii'jkt}}{R_{ii'j'kt}} = \left(\dfrac{X_{ijkt}}{X_{i'jkt}}\right) \Big/ \left(\dfrac{X_{ij'kt}}{X_{i'j'kt}}\right) = \left(\dfrac{\phi_{ijkt}}{\phi_{ij'kt}}\right) \Big/ \left(\dfrac{\phi_{i'jkt}}{\phi_{i'j'kt}}\right).$

The *RORs* only depend upon purely bilateral trade determinants. Clarifying how these determinants relate to tariffs requires being more specific about demand. A standard assumption, consistent with the general equations presented above is that the elasticity of substitution between imports from different origins is constant. In other words, the bundle of imports for a given product can be represented within consumer preferences through a constant elasticity of substitution (CES) function over imports of different origins. Practically, country j's consumption sub-index specific to good k imports at period t is assumed to write:

(2) $C_{jkt} = C_{jkt_0} \left[\displaystyle\sum_{i \neq j} \theta_{ijk} \left(\dfrac{q_{ijkt}}{q_{ijkt_0}} \right)^{\frac{\sigma-1}{\sigma}} \right]^{\frac{\sigma}{\sigma-1}},$

where q_{ijkt} refers to the quantity of good k originating in country i consumed in country j at year t, t_0 denotes a year of reference, $\sigma > 0$ designates the elasticity of substitution between varieties of good k. Given the calibrated share form used, $\theta_{ijk} = \dfrac{p_{ijkt_0} q_{ijkt_0}}{\sum_{l \neq j} p_{ljkt_0} q_{ljkt_0}}$, where p_{ijkt} refers to the tax-inclusive price of country j imports of products k from country i at year t. θ_i is thus the value share of supplier i in country j imports of good k at year t_0. The corresponding dual price index is:

(3) $P_{jkt} = P_{jkt_0} \left[\sum_{i \neq j} \theta_{ijk} \left(\dfrac{p_{ijkt}}{p_{ijkt_0}} \right)^{1-\sigma} \right]^{\frac{1}{1-\sigma}}.$

Import demand in (tax-inclusive) value can be expressed as (4) and under mill pricing and assuming that bilateral transport costs (as well as regulations with possible impact on prices) do not change over time, tariffs are the only source of change in prices, hence (5) holds:

(4) $p_{ijkt} q_{ijkt} = p_{ijkt_0} q_{ijkt_0} \left(\dfrac{p_{ijkt}}{p_{ijkt_0}} \right)^{1-\sigma} \left(\dfrac{P_{jkt_0}}{P_{jkt}} \right)^{-\sigma} \left(\dfrac{C_{jkt}}{C_{jkt_0}} \right),$

$X_{ijkt} = X_{ijkt_0} \left(\dfrac{\tau_{ijkt}}{\tau_{ijkt_0}} \right)^{-\sigma} \left(\dfrac{P_{jkt_0}}{P_{jkt}} \right)^{-\sigma} \left(\dfrac{C_{jkt}}{C_{jkt_0}} \right),$

where, as before, X refers to *tax-exclusive* import values. $\tau_{ijkt} - 1$ is the *ad valorem* customs duty applied by country j to imports of good k from supplier i at time t. It is noteworthy that the same equation would have been obtained, had CES preferences been defined over a basket including domestic consumption of good k, in addition to the imports. This equation clarifies the link between applied tariffs and bilateral determinants of trade flows. The *ROR* defined in (1) then writes

(6) $\quad ROR_{ii'jj'kt} = ROR_{ii'jj'kt_0} \left[\left(\frac{\tau_{ijkt_0}}{\tau_{i'jkt_0}} \right) \Big/ \left(\frac{\tau_{ij'kt_0}}{\tau_{i'j'kt_0}} \right) \right]^{\sigma} \left[\left(\frac{\tau_{ijkt}}{\tau_{i'jkt}} \right) \Big/ \left(\frac{\tau_{ij'kt}}{\tau_{i'j'kt}} \right) \right]^{-\sigma}$,

Now, let the indices i and j represent two partners having signed a bilateral trade agreement. Also, let j' denote a control market, defined for each importer $[j' = J(i)$, referred to as J in what follows for the sake of brevity] as a representative set of countries whose trade policy regarding country i has not changed during the period under study $[\tau_{iJkt} = \tau_{iJkt_0}]$.[3] Finally, let i' be a control group of exporters $[i' = I(j)$, henceforth $I]$ consisting of trading partners such that the trade policy of both country j and the control market J toward this group I remained unchanged during the period under examination $[\tau_{iJkt} = \tau_{iJkt_0}, \tau_{IJkt} = \tau_{IJkt_0}]$.[4]

Under these conditions, the ratio of import duties applied by the reference group of importers J to suppliers i and I does not change over time. Equation (6) can then be rewritten as $ROR_{iIjJkt} = \lambda_{ijk} \left(\frac{\tau_{ijkt}}{\tau_{Ijkt}} \right)^{-\sigma}$ where $\lambda_{ijk} = ROR_{ii'jj'kt_0} \left[(\tau_{ijkt_0} / \tau_{i'jkt_0}) \right]^{\sigma}$, I and J are respectively functions of j and i, so that *ROR* only depends on i, j and k. Traditionally, this type of equation has been estimated in log-linear form:

(7) $\quad \ln(ROR_{ijkt}) = \alpha_{ijk} - \sigma \ln \left(\frac{\tau_{ijkt}}{\tau_{Ijkt}} \right) + u_{ijkt}$,

where u represents an error term and $\alpha_{ijk} = \ln(\lambda_{ijk})$. This specification includes one fixed effect specific to each exporter-importer-good triplet. As a consequence, the elasticity of substitution between imports from different origins σ is only estimated out of changes over time within each of these triplets. Fixed effects by exporter, importer, or good, or by any combination of two of these dimensions, are implicitly accounted for.[5]

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[3] When practically building the control groups, we interpret this condition of unchanged trade policy as meaning that no RTA has been signed or phased in between the countries.

[4] A sufficient condition is actually that the ratio τ_{iJkt}/τ_{IJkt} remains constant over time. Here again, we interpret this condition of unchanged trade policy as meaning that no RTA has been signed or phased in between the countries.

[5] It is superfluous to incorporate them explicitly since they would be perfectly correlated with the fixed effects already included. Two-way time-exporter or time-importer fixed effects are not needed either, since the corresponding shocks should be absorbed by the dependent variable transformation.

Santos-Silva and Tenreyro (hereafter SS&T) demonstrate the bias inherent in estimating such models in their logarithmic form (SS&T 2006). Beyond the failure to account for null flows, they show that this inconsistency mainly results from the logarithmic transformation of the empirical model, in a context of marked heteroskedasticity. Both of these concerns are likely to be aggravated when estimates are carried out at the product level, where zeroes are widespread and magnitudes are highly variable. SS&T (2006) argue that the gravity equation should be estimated in its multiplicative form, and suggest using a Poisson Pseudo-Maximum Likelihood (PPML) estimation technique. These arguments, developed for a standard gravity model, also apply to the present case, which is a difference-in-differences of the log-transformed model, or a ratio of ratios of the multiplicative form. In particular, no distributional assumption is needed to ensure consistency of the PPML estimator, which only requires correct specification of the conditional mean.

SS&T (2011) show the theoretical consistency and the good statistical properties of this estimator and how it performs better than proposed alternatives (see also Fally 2015; Sun and Reed 2010). They also show Gamma PML to be an efficient estimator in this context. As discussed in Head and Mayer (2014), though, these results are based upon data generating process which do not reflect modern theories of international trade. Ultimately, the best estimator depends upon the form of the heteroskedasticity, since PPML is efficient under the assumption of constant variance to mean ratio, while Gamma PML is efficient under constant coefficient of variation assumption. The Manning and Mullahy (2001) test (referred to as "MaMu" test in Head and Mayer, 2014, and as Park-type test in SS&T 2006) allows the corresponding diagnostic to be made, by studying the relationship between the empirical counterparts of the error variance and of the dependant variable's expected value.[6] While providing a biased estimate of the corresponding coefficient, Head and Mayer (2014) show that this test usefully discriminates between alternative settings. Here, applying this test to the benchmark estimation presented below (estimate 1 in table 2) delivers an estimated coefficient $\hat{\lambda} = 1.46$, with a standard error of 0.31. Accordingly, our preferred estimator in what follows is the PPML.[7]

Our estimates are thus based upon fixed-effect PPML estimates of the multiplicative form of the model (8):

$$(8) \quad ROR_{ijkt} = \exp\left(\alpha_{ijk} - \sigma \ln\left(\frac{\tau_{ijkt}}{\tau_{Ijkt}}\right)\right) + v_{ijkt}.$$

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[6] Practically, the test is based upon a regression where the dependent variable is the logarithm of the squared error term, the latter being computing as the difference between trade in level and the exponential of its fitted log-value. The fitted log-trade value is used as independent variable.

[7] The PPM estimator is efficient for $\lambda = 1$, while Gamma PML is efficient for $\lambda = 2$, but Head and Mayer (2014) show that OLS estimates of λ such as the ones used here are significantly biased upward, so that "estimates of λ significantly below two were a near perfect predictor of a constant variance to mean ratio" in their simulations.

The dependent variable computation involves control groups' trade flows: for a given importer, these flows are common to exporters; for a given exporter, they are common to importers. This feature might originate residual correlation across observations for a given exporter or a given importer, in a given year. For this reason, standard errors should ideally be clustered by both importer-year and exporter-year. This is possible for OLS estimates, based on Cameron, Gelbach and Miller (2011), but not for PPML estimates. In the latter case, we compute robust standard errors clustered at the panel level (i.e., by exporter-importer-product triplet), following Wooldridge (1999). However, in order to be able to compute clustered standard errors, Ordinary Least Square (OLS) estimates of the log-transformed model based on (7) are carried out in addition.

As the recent international trade literature has clearly demonstrated, it is important to account, not only for trends in existing trade flows, the so-called intensive margin of trade, but also for developments in the number of goods traded at the extensive margin- (e.g., Chaney 2008 or Helpman, Melitz and Rubinstein 2008). The estimates described so far focused on the intensive margin (defined at the country-pair, product level), since they deal with changes over time in non-zero trade flows. As a complement, we estimate the probability of exporting. Econometric modeling of this probability requires accounting for determinants specific to each (potential) exporter-importer-good triplet. Exporter-by-year and importer-by-year fixed effects are also warranted here, since the corresponding idiosyncratic shocks are not controlled by the transformation of the dependent variable, as they were above. Given the large size of the sample, accounting for fixed effects by panel unit is only possible using a Within estimator—explicitly incorporating all the dummies is numerically intractable.

Following Frazer and Van Biesebroeck (2010) and Head, Mayer and Ries (2010), we use a linear model to estimate the probability of exporting. The estimating equation is:

$$(9) \quad P\big(X_{ijkt} > 0\big) = \alpha'_{ijk} + \beta'_{it} + \gamma'_{jt} - \sigma' \ln\left(\frac{\tau_{ijkt}}{\tau_{Ijkt}}\right) + w_{ijkt},$$

where *w* is an error term. Our benchmark estimates thus include, in addition to the panel-specific fixed effects α'_{ijk} (81,570 in the benchmark estimates), 419 reporter-year fixed effects (35 countries x 12 years, minus one) and as many partner-year fixed effects. As argued by Frazer and Van Biesebrock, "the main disadvantage [of the linear probability model], that predicted values are not restricted to lie on the (0,1) interval, is unlikely to be much of an issue as all coefficients are identified off the time variation within country-product categories." In addition, Angrist and Pischke (2009:107) emphasize that linear estimates, while theoretically less well-suited, are very close in practice to the marginal effects drawn from non-linear models.

2. Data

The method presented above requires panel data on bilateral trade and preferential tariffs, at the product level. We first relied upon a study jointly undertaken by the OECD and the Inter American Development Bank on the treatment of agriculture in trade agreements, in order to characterize the concession schedules, product by product (Fulponi, Shearer and Almeida 2011). Once duly codified, this information was combined with assessments of ad-valorem equivalent of MFN tariff duties, taken from the MacMap-HS6 database for years 2001, 2004 and 2007, and encompassing both ad-valorem and non-ad valorem duties (see Guimbard et al. 2012). This made possible the creation of a database giving ad-valorem equivalents of preferential duties for 74 RTAs, over the period 1998-2009. Trade data (cost-insurance-freight, inclusive of annual imports) are sourced from the BACI database, which provides a detailed description of the role played by RTAs in the global trade of agricultural products over the 1998-2009 period (Gaulier and Zignago 2010). However, composition of the sample RTAs under study is constrained by data availability and cannot be deemed representative. In particular, Latin-American countries are overrepresented, while few African RTAs are represented. A broader application of the methodology would be desirable. Still, the large size of the sample already makes it possible to deliver meaningful insights for RTAs in general.

Application of the ratio of ratios method requires the creation of control groups specific to each importer and exporter. For a given country, the corresponding control group consists of all countries in our database that had not signed a preferential agreement with this country by 2009. For the sake of robustness, this group includes the entire population of countries that are meaningful points of comparison. This composition is specific to each country, but stable over time. Estimates are run at six-digit product code level of the Harmonized System.

Working with product-level trade data entails a risk of lack of robustness, raising a particular problem when the dependent variable is computed as a ratio as is the case here. We deal with this concern by using several checks. First, to make sure that trade flows with control groups are representative, we only retain those exceeding USD 200,000.[8] Second, we drop observations where either of the control ratios X_{ijkt}/X_{IJkt} and X_{Ijkt}/X_{IJkt} deviates from its median over the period by a multiplicative factor of more than three:[9] when the deviation exceeds this factor, we consider the ratio's instability to prevent it from being a reliable

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[8] This condition only applies to trade with control groups, not to trade between the partner and the reporter. The threshold has been chosen based on an analysis of the degree of autocorrelation of product-level trade flows. Indeed, control groups' trade flows are used as benchmark, representative of partner-specific trade determinants. Since these determinants are likely to be highly correlated in practice (they are linked to variables such as preferences and productivity), a low degree of autocorrelation denotes a lack of representativeness, which is likely when flows are small, both because they may depend excessively upon incidental factors, and because of the interference with the statistical reporting thresholds used by many countries. As a matter of fact, flows autocorrelation is significantly lower when their magnitude is below USD 200,000. Robustness checks were carried out with alternative thresholds of USD 100,000 and USD 500,000 (details on autocorrelation and robustness checks are available in the Appendix).

[9] Robustness checks based on a multiplicative factor of four or two instead of three are presented in the Appendix.

control. Third, robustness checks are carried out, whereby estimates are limited to partner-reporter-product triplets for which the average annual trade is more than USD 100,000.

Another concern is the measurement of protection during the first year of enforcement of an RTA. Indeed, protection is measured in our database on January 1 of each year. For an RTA enforced over the course of the year, tariff cuts will thus only be taken into account as of the following year, potentially resulting in serious mismeasurement of applied tariffs on the first year. In addition, the trade impacts of an RTA may be somewhat delayed with respect to enforcement, due to the need for involved economic agents to adapt to a new institutional context. We consider separately the impact of RTAs on the first year of entry into force. Practically, the term measuring preferential treatment $\left(\tau_{ijkt}/\tau_{Ijkt}\right)$ is a dummy indicating whether the RTA between i and j, if any, is in its first year of implementation.

3.   RTAs and International Trade Flows

The share of global trade conducted between two partners having signed a trade agreement rose from less than 24% in 1998 to over 36% in 2009 (the year in which our data sample ends), without showing any sign of slowing at the end of the period. An analysis by broad sector reveals that while the share of global trade occurring between RTA partners was higher for manufactured products in 1998, it fell below that of agro-food products by 2009. The accelerated pace recorded by the agricultural sector might at least partly reflect the greater intensity of agricultural trade between signatories to RTAs that entered into force between 1998 and 2009.

The implementation of these agreements typically stretches out over some 10 years, and the transition period occasionally exceeds 15 years for selected products. To assess the corresponding concessions, we compute "preferential margins" defined as the price wedge (taxes included) attributable to preferential treatment. Note that this is different from the reduction in tariff rates as the one appearing in equation (7). For *ad valorem* import duties under the MFN system (t_{MFN}) and under the preferential system (t_{pref}), this margin is thus defined as $m = 1 - \left(1 + t_{pref}\right)/\left(1 + t_{MFN}\right)$ or, in terms of our earlier notation, $m = 1 - \tau_{pref}/\tau_{MFN}$.

A detailed examination of the tariff concessions for agricultural and food products in the 74 RTAs allows the average preference margin to be computed for each agreement as a function of the number of years elapsed since entry into force, based on the exact phase-in pattern. The mean preferential margin nearly doubles within eight years of its entry into force, rising from 4.3% during the first year to 8.8% eight years later. In contrast, the level of preferential margins hardly changes beyond 10 years of entry into force of an agreement.

Distinguishing between two groups of partner countries: high-income members of the OECD (referred to as the "North" in this article) and the others (the "South") reveals that preferential margins granted by agreements between countries of the South at the end of the phase-in

period (10.1%) are near the mean calculated for all agreements (9.5%). Conversely, North-South agreements are asymmetric, especially after several years: margins granted by countries of the South are higher (10.4% after 8 years, 11.2% once fully phased in) than those granted by countries of the North (6%, with little variation over time).

A similar calculation demonstrates the differences between the chapters of the Harmonized Commodity Description and Coding System (HS), in Table 1. Significant variation is evidenced, since the average preferential margin reached at the end of the phase-in period ranges from 5.4% for non-food agricultural products to as high as 14.0% for dairy products, eggs and honey (Chapter 2) and 16.8% for tobacco products (Chapter 24). The preferential margin exceeds 10 points in more than half agricultural sectors. The cross-sectoral differences in preferential margin tend to increase over time elapsed since implementation, illustrating the well-known fact that transition period are usually longer for more sensitive products.

Table 1. Mean base rate and preferential margin by Harmonized System chapter and by time elapsed since entry into force of the agreement (in percent)

Chapter	Base rate	Preferential margin			Chapter	Base rate	Preferential margin		
		Year 1	Year 5	Full			Year 1	Year 5	Full
01- LIVE ANIMALS	9.4	5.3	6.0	6.5	14- VEGETABLE PLAITING MATERIALS	6.0	3.6	5.1	5.9
02- MEAT & EDIBLE MEAT OFFAL	21.8	5.1	8.4	11.8	15- ANIMAL OR VEGETABLE FATS & OILS	10.8	3.7	6.4	8.9
04- DAIRY PRODUCE; EGGS; HONEY	27.9	7.0	10.4	14.0	16- PREPARATIONS OF MEAT & FISH	20.1	4.9	9.0	12.7
05- PROD. OF ANIMAL ORIGIN, NES	5.8	4.1	5.0	5.6	17- SUGARS & SUGAR CONFECTIONERY	18.3	4.9	7.7	10.9
06- LIVE TREES & OTHER PLANTS	10.1	5.2	6.8	7.6	18- COCOA & COCOA PREPARATIONS	11.2	4.9	7.4	9.8
07- VEGETABLES	13.3	5.9	8.8	10.5	19- PREP. OF CEREALS	13.6	4.8	8.5	11.4
08- FRUITS	11.6	6.3	8.8	10.4	20- PREP. OF VEGETABLES & FRUITS	15.0	6.1	10.0	12.7
09- COFFEE, TEA, SPICES	10.0	4.9	7.1	8.4	21- MISCELLANEOUS EDIBLE PREP.	13.5	6.0	9.5	12.0
10- CEREALS	17.1	5.6	7.7	10.1	22- BEVERAGES, SPIRITS & VINEGAR	23.6	6.1	10.5	13.5
11- PROD. OF THE MILLING IND.	17.1	4.4	8.4	11.6	23- FOOD RESIDUES & WASTE	9.3	4.1	6.1	7.6
12- OIL SEEDS & OLEAG. FRUITS	7.0	4.0	5.2	5.9	24- TOBACCO	23.4	9.5	13.6	16.8
13- LAC; GUMS, RESINS	5.9	3.5	5.0	5.7	NON-FOOD AG. PRODUCTS	5.9	3.2	4.5	5.4
					All products	13.0	5.0	7.5	9.5

Source: Calculated by the authors from the BACI (CEPII) database, Comtrade (UN), MAcMap-HS6, and IDB data.

Note: "Year 1" refers to the year following the entry into force of the agreement, "Year 5" to the fifth year after entry into force, "Full" to the full implementation of RTAs, once the phase-in period is over. "Base rate" refer to the duty rate used as a basis for the agreement. The mean si computed over all agreements covered in the paper (see list in Appendix).

Illustrating how RTAs may be related to trade flows is difficult because of marked cross-country differences in import growth trend. Focusing on the share each signing country represents in its RTA partner's imports allows to get rid of these influences. As shown in Figure 1, on average, this share increases slightly (by less than 10% after three years, compared to the year preceding enforcement) and with a delay following an RTA enforcement when all products are considered jointly. For agricultural and food products, the

increase is more pronounced as of the year of entry into force of the RTA, with an order of magnitude of 15 to 20% after 3 years.

Figure 1. RTAs and share in partner's imports
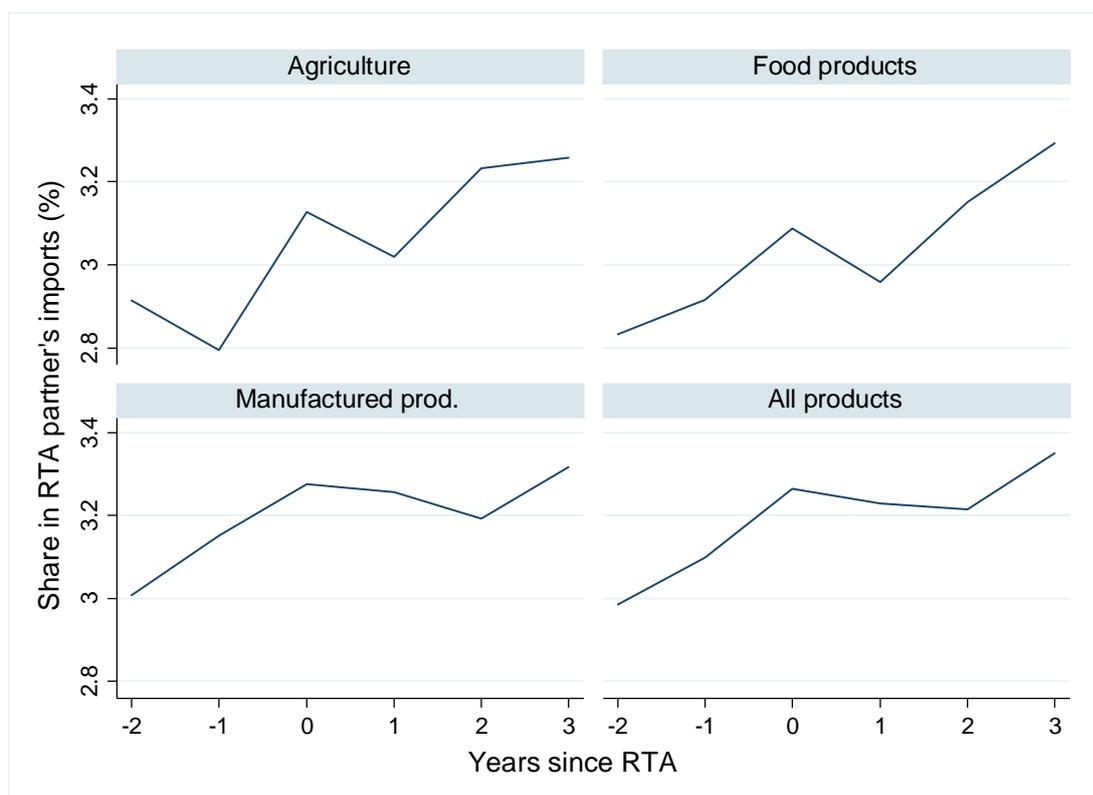
Source: Calculated by the authors from the BACI (CEPII) database, Comtrade (UN), MAcMap-HS6, and IDB data. Scope: 384 ordered country pairs among for which an RTA entered into force between 2000 and 2006.

The impact of RTAs on trade may be blurred by cross-country differences in export growth trends, though. To control for this possible influence, Figure 2 represents the share of the RTA partner in a country's imports as a proportion of the share of all non-RTA partners in the country's imports. Using this metrics provides a more telling picture of the trade impact of RTA. On average across all products, this relative import share of RTA partners is increased by 20% three years after enforcement; for agricultural and food products, it is increased by 35% to 40%.

Figure 2. RTAs and share in partner's imports, relative to non-RTA partners
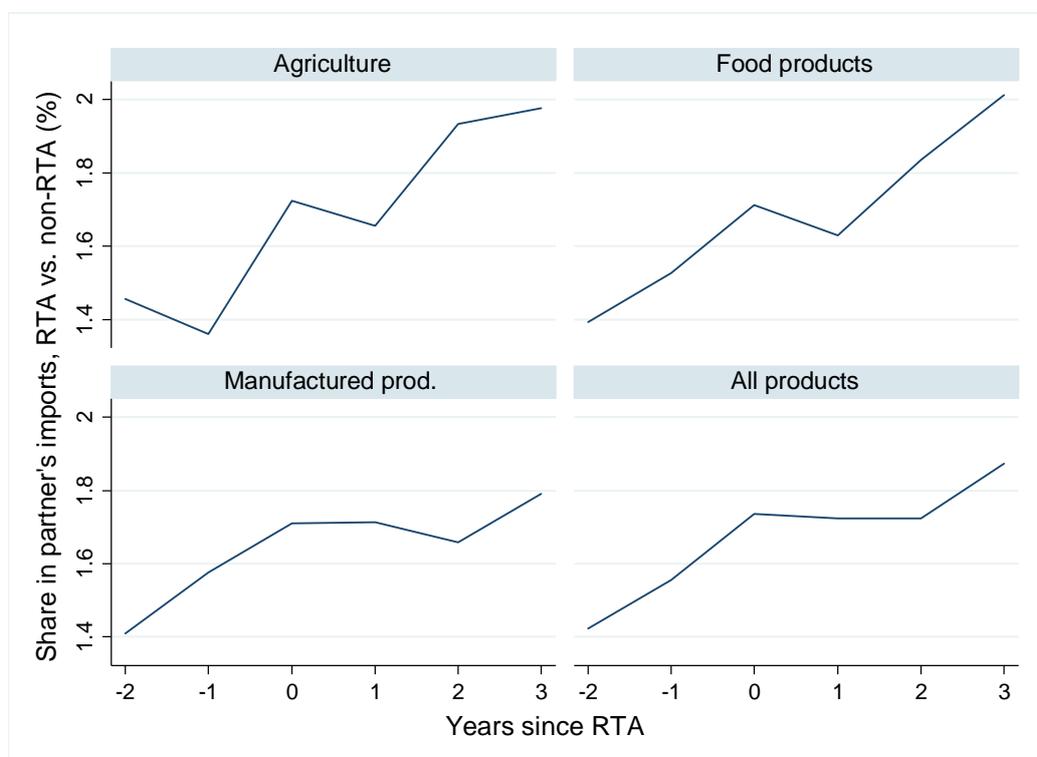
Source: Calculated by the authors from the BACI (CEPII) database, Comtrade (UN), MAcMap-HS6, and IDB data. Scope: 384 ordered country pairs among for which an RTA entered into force between 2000 and 2006.

Note: The relative share is computed as the share of the RTA partner in the importer's total imports, divided by the share of non-RTA partners in the importer's total imports.

4. Estimation results

To evaluate the impact of RTAs on international trade in agricultural products, the method described above is jointly applied to the 74 agreements for which we have been able to obtain complete information on both the ad valorem equivalents of customs duties and the concession schedules, over the 1998-2009 period.

4.1. Intensive margin

Our benchmark estimate of the import substitution elasticity, obtained by applying the PPML estimator to the whole sample, gives a statistically significant value of 3.71 at least one year after implementation (estimate 1 row 3, Table 2). This means that an RTA cutting the applied tariff by 1% increases bilateral exports by 3.78% (exp(3.71)-1 = 3.78%), compared to flows between signatory parties and the control groups. This increase actually concerns the ratio of

bilateral exports to its controls, referred to above as bi-ratios; as such, it may results from an increase in bilateral exports (trade creation), but also from a decline in trade with third countries (trade diversion). Although our methodology does not allow the two effects to be disentangled, the conclusions that may be drawn are clarified in the counterfactual simulations below.

The specificity of the first year of implementation of RTAs, presumably linked to a combination of measurement issues and adjustment delays, is confirmed because the corresponding estimated elasticity (2.22) is about half as large as it is afterwards and only statistically significant at the 10% level.[10] If anything, this confirms that an analysis of the year of implementation is difficult based on annual data.

The base OLS estimate of the import-substitution elasticity (after the first year of implementation of preferential agreements) is also statistically significant, but far lower than the PPML one (0.94, see estimate 2, table 2). OLS estimates are known to be biased in this context, but the method allows clustered standard errors to be computed, which in this case remain rather small, corresponding to a significance level comparable to the one found for PPML estimates. As a robustness check, the estimates are also carried out on a sample which excludes panels (i.e., reporter-partner-product triplets) for which the average of non-zero trade flows over the 1998-2009 period is lower than USD 100,000. Indirectly, this is a way to control for the possible lack of reliability of statistics when trade flows are low. These estimates are referred to as "Excluding low mean panels" in the following tables. Such restriction does not change the PPML estimate much (3.92, see estimate 3), but it doubles the OLS estimate (1.70), suggesting that the latter is more sensitive to small-flow observations. As an additional robustness check, the sample is further restricted to the 2001-2009 period because information on MFN duties is less reliable before 2001, the first year it became directly available (estimates referred to as "Excluding low mean panels post 2000" in the following tables). Doing so does not substantially change the value of the PPML estimate (3.54, see estimate 5, table 2), but the OLS estimate is increased to 2.20. Unreported robustness checks were carried out restricting the control group to countries within the same continent or within the same income category; alternatively, regressions were weighted by the log of the value of trade flows.

On the whole, PPML estimates appear consistent and robust, which is not the case of OLS estimates. A possible explanation is that the inconsistency of OLS estimates, which are based on the log transformation of the model, is aggravated when analysis is carried out at the product level, since magnitudes are even more variable than across countries.[11] This

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[10] With exceptions, estimated elasticities for the first year of implementation, not commented in the article and not shown in tables, are small and insignificant statistically.

[11] Zero flows are more widespread, but this is not fully apparent here, since the dependent variable is also frequently missing, due to zeroes in control groups' trade flows (or to values too low to be considered representative, as mentioned above).

seems to matter even when identification is only achieved out of time variations within reporter-partner-product triplets.

Table 2. Estimated import substitution elasticity for agricultural products

	Benchmark estimates		Excluding "low-mean" panels		Excluding "low-mean" panels, post-2000	
	(1)	(2)	(3)	(4)	(5)	(6)
FTA's first year	2.22 *	0.08	2.33 *	0.69 *	2.17	0.98 **
	(1.27)	(0.36)	(1.38)	(0.42)	(1.37)	(0.43)
Subsequent years	3.71 ***	0.94 ***	3.92 ***	1.70 ***	3.54 ***	2.20 ***
	(1.16)	(0.28)	(1.25)	(0.34)	(1.14)	(0.37)
Estimation method	PPML	OLS	PPML	OLS	PPML	OLS
Observations	80,641	63,122	47,127	43,128	36,658	33,876
Panel units	10,703	9,462	5,888	5,678	5,624	5,433

Source: Authors' estimates based on IDB, MAcMap-HS6 (ITC and CEPII), BACI (CEPII) and Comtrade (UN Statistics Division) data.

Notes: All estimates include three-way partner-reporter-product fixed effects. Estimates are based on equation (8) for PPML estimates and on equation (7) for OLS estimates. The coefficients reported refer to the estimated value of σ, interacted with a dummy indicating whether the FTA is in its first year of implementation (first row) or not (second row). "Low-mean panels" are those partner-reporter-product triplets for which the mean value of non-zero trade flows is lower than USD 100,000 (see text). The estimation period is 1998-2009, except in columns (5) and (6), for which it is restricted to 2001-2009. Robust standard errors are reported in parenthesis, clustered at the panel (reporter-partner-product) level for PPML estimates, at both reporter-year and partner-year level for OLS estimates. Asterisks denote statistical significance levels 10% (*), 5% (**) and 1% (***) respectively.

Romalis (2007) is a natural point of comparison for these results, since our methodology is partly based on that paper. His estimates vary between 6.3 and 9.4 for U.S. imports from Canada, between 9.6 and 10.9 for U.S imports from Mexico, between 2.8 and 5.5 for Canadian imports from the United States, between 6.6 and 8.1 for Canadian imports from Mexico, between 2.0 and 2.5 for Mexican imports from the United States, and between 0.5 and 0.7 (not significant) for Mexican imports from Canada. Our estimates seem consistent with these figures, though somewhat lower. This impression is borne out by other detailed estimates that are available, even though none of them is specific to agriculture. Working with a very different methodology, Broda and Weinstein (2006) find that the un-weighted mean of elasticities of substitution estimated for the United States between 1990 and 2001 is approximately 12.6 for goods at the ten-digit product code level of the HS (in comparison to only 4.0 for the three-digit product code level), for a median of 3.1 (2.2 at three digits). Simonovska and Waugh (2014) find approximately 4 for the same product code level as the one we use, HS6. Estimated import-demand elasticities (at HS6 level) by Kee, Nicita and Olarreaga (2008) equal 3.1 on average for all products. Head and Mayer's (2014) meta-

analysis of estimates based on aggregate date shows a median of 5 for structural estimates using tariffs or freight rates as identifying variable. All in all, these comparisons suggest that our estimates are in line with others in the literature, although perhaps on the low side.

To investigate whether these effects depend on the income level of partners, the agreements are again grouped according to the categorization of partner countries into North and South. The estimated elasticity is larger for North-to-South flows than for South-to-North and South-South flows in each case. While statistically significant,[12] the differences between estimated elasticities remain of limited extent (respectively 4.55, 3.74 and 2.97 in the benchmark estimation, see table 3, estimates 1-3).

Table 3. Estimated import substitution elasticity for agricultural products, by type of agreement, by level of preference margin and by HS section

	Benchmark estimates			Excluding "low-mean" panels			Excluding "low-mean" panels, post-2000		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
South-South	3.74 **			3.95 **			3.56 *		
	(1.80)			(1.92)			(1.99)		
North-to-South	4.55 ***			5.14 ***			3.99 **		
	(1.33)			(1.58)			(1.71)		
South-to-North	2.97 ***			2.98 ***			3.61 ***		
	(1.03)			(1.05)			(1.05)		
0 < preferential margin < 5%		5.12			5.85			5.99	
		(3.46)			(3.69)			(3.83)	
5% < preferential margin < 10%		3.15 *			3.48 *			5.05 ***	
		(1.68)			(1.79)			(1.86)	
10% < preferential margin		3.81 ***			4.02 ***			3.29 ***	
		(1.19)			(1.28)			(1.00)	
I Animals & products			3.66			3.11			0.34
			(2.45)			(2.49)			(4.37)
II Vegetable products			-0.34			-0.80			-0.71
			(2.07)			(2.33)			(2.72)
III Fats & oils			8.88 **			9.89 ***			8.32 *
			(3.60)			(3.78)			(4.41)
IV Foodstuffs, bev., tobacco			4.09 **			4.47 ***			4.76 ***
			(1.62)			(1.73)			(1.47)
Non food ag. products			8.93 **			9.47 **			5.09
			(4.07)			(4.27)			(4.65)
Observations	80,641	80,641	80,641	47,127	47,127	47,127	36,658	36,658	36,658
Panel units	10,703	10,703	10,703	5,888	5,888	5,888	5,624	5,624	5,624

Source: Authors' estimates based on IDB, MAcMap-HS6, BACI and Comtrade data.

Notes: All results are PPML estimates based on equation (8). The coefficients reported refer to the estimated value of σ in cases where the FTA is beyond its first year of implementation. All estimates include three-way partner-reporter-product fixed effects. The estimation period is 1998-2009, except in columns (7), (8) and (9), which are restricted to 2001-2009.

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[12] A likelihood-ratio test rejects the equality of the corresponding coefficients (χ^2= 12.3, p-value < 1% in the benchmark estimation).

Another question is the stability of elasticities over different ranges of preferential margins (recall that the term "preferential margin" is used here to designate the price wedge, taxes included, attributable to preferential treatment). Consider separately cases where the preferential margin is lower than 5%, between 5% and 10%, and higher than 10% yield differences which, although statistically significant,[13] are not substantial. The estimates are insignificant in the first category, perhaps by lack of variance, and less precise in the intermediate one, but differences again remain proportionately limited. Elasticities may also differ across products. Identification of the model would be problematic with product-specific elasticities, by lack of variance. Instead, we assume elasticities to be constant across HS sections. Doing so, we find high elasticities for fats and oils (11.1 in the base case) and for non-food agricultural products (8.05), while estimated elasticities for both animals and their products and vegetable products are rather low, although not precisely identified.[14]

4.2. Extensive margin

To assess the incidence of preferential tariffs on the extensive margin, our estimates make use of a linear model of the probability of exporting, as described above. In this case, our sample is nearly 10 times larger than before, because there is no more need to use representative control-group trade flows. Our estimates confirm the impact of preferential tariffs on the probability of exporting. After the first year (for which the unreported estimated effect is insignificant), the estimated effect across all agreements is relatively weak (0.045), suggesting that a preferential margin that lowers the tariff-inclusive price by 10% increases the probability of exporting by 0.45% (table 4, estimation 1).

Distinguishing across categories of agreements shows that this effect is only significant for North-to-South trade flows, for which the corresponding elasticity (0.13) is thrice as large as the one found on average (table 4, estimation 2). Worth noting, this category is the one for which the estimated intensive-margin elasticity was the lowest. Distinguishing by level of preferential margin also shows that the elasticity is higher for low preferential margin levels, suggesting that the potential for trade preference to create new trade flows does not increase much beyond a certain level of preferential margin.

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[13] A likelihood-ratio test rejects the equality of the corresponding coefficients ($\chi^2 = 34.0$, p-value < 1% in the benchmark estimation).

[14] The equality of coefficients across sections is rejected by a likelihood-ratio test ($\chi^2 = 183.3$, p-value < 1% in the benchmark estimation).

Table 4. Estimated impact of preferential duties on the export probability

	Dependent variable: Export probability		
	(1)	(2)	(3)
Independent variable:			
Log price wedge linked to preferential duties			
All agreements, all products	0.045 ***		
	(0.014)		
South-South		0.009	
		(0.017)	
North-to-South		0.127 ***	
		(0.033)	
South-to-North		0.040	
		(0.033)	
Elasticity, Pref. margin 0-5%			0.108 **
			(0.045)
Elasticity, Pref. margin 5-10%			0.116 ***
			(0.024)
Elasticity, Pref. margin > 10%			0.041 ***
			(0.014)
Observations	978,840	978,840	978,840
Panel units	81,570	81,570	81,570

Source: Calculated by the authors from IDB, MAcMap-HS6, BACI and Comtrade data.

Notes: All estimates use a linear probability model, based on equation (9). The coefficients reported refer to the estimated value of σ', in cases where the FTA is past its first year of implementation. All estimates include three-way partner-reporter-product fixed effects, as well as two-way partner-year and reporter-year fixed effects. The estimation period is 1998-2009. Robust standard errors, clustered at the panel (reporter-partner-product) level, are reported in parenthesis.

4.3. Preferential margin categories

A complementary approach for evaluating the impact of preferential treatment consists of grouping exporter-importer-good triplets into several categories, reflecting whether an agreement is in effect and, if so, the magnitude of the preferential margin. To do this, we create five categories: (i) no agreement is in place; (ii) an agreement is in force but the preferential margin is zero for this product; (iii) the preferential margin is greater than zero, but less than 5%; (iv) the preferential margin is between 5 and 10%; and (v) the preferential margin is greater than 10%. In the econometric specification defined by equation (8), the level of tariffs is replaced by a series of dummies indicating belonging to one of these categories, with the first group serving as the reference. As before, the first year of the implementation of each agreement is handled separately. For the sake of clarity, only the coefficients of the effects during subsequent years are presented.

Estimates indicate that an agreement does not materially affect trade creation for products that do not benefit from a preferential margin (table 5, estimates 1-3, first line). The trade impact is not found to be significant either when the preferential margin is lower than 5%. Conversely, according to our results, entry into force of a trade agreement translates into a strong and significant increase in bilateral exports relative to third countries in the case of good benefitting from a preferential margin larger than 5%.

According to our benchmark estimates, the increase averages approximately 27% [exp(0.24)–1=27%] when this margin is between 5% and 10%, and 60% when the preferential margin exceeds 10%. Applying the same approach to the export probability only gives statistically significant impacts for products where the margin is between 5% and 10%, in which case the probability is found to be increased by 0.8% (table 5, estimation 4). This confirms the limited impact of RTAs on the extensive margin for agricultural and food products.

Table 5. Estimation of the trade impact of preferential agreements, by level of preferential margin

Dependant variable:	Diff in diff log exports			Export probability
		Excl. low-	Excl. low,	
Sample:	All	mean panels	post-2000	All
	(1)	(2)	(3)	(4)
Independent variable:				
Dummy variable indicating				
PTA in force,				
preferential margin = 0 for this product	0.00	0.02	0.08	-0.004
	(0.11)	(0.13)	(0.13)	(0.003)
0 < preferential margin < 5%	0.16	0.19	0.21 *	0.003
	(0.12)	(0.13)	(0.12)	(0.002)
5% < preferential margin < 10%	0.24 *	0.27 *	0.41 **	0.008 ***
	(0.14)	(0.15)	(0.16)	(0.003)
10% < preferential margin	0.47 ***	0.51 ***	0.36 **	0.004
	(0.17)	(0.18)	(0.15)	(0.003)
Estimation method	PPML	PPML	PPML	Linear prob. model
Observations	80,641	47,127	41,670	978,840
Panel units	10,703	5,888	6,346	81,570

Source: Authors' estimates based on IDB, MacMap-HS6, BACI and Comtrade data. Notes: Estimates (1)-(3) are based on equation (8), column (4) is based on equation (9). The coefficients reported refer to the estimated value of σ and σ', respectively, in cases where the FTA is beyond its first year of implementation. In each case, the tariff duty variable is replaced by a set of dummy variables, as indicated in the text. All estimates include three-way partner-reporter-product fixed effects. Estimate (4) also includes two-way partner-year and reporter-year fixed effects. The estimation period is 1998-2009, except for (3), where it is 2001-2009. Robust standard errors, clustered at the panel (reporter-partner-product) level, are reported in parenthesis.

5.   Counterfactual simulations of the trade impacts of RTAs

The estimates above allow the partial equilibrium trade impacts of each RTA to be simulated, product by product. This section lays out the corresponding theoretical underpinnings and briefly presents the results obtained.

5.1.   Underpinnings and assumptions

A fully consistent assessment of the impact of an RTA on trade flows would require a general-equilibrium analysis taking into account, *inter alia*, supply responses and changes in import demand. Here, we make two simplifying assumptions. First, supply response is not considered. In most cases, the RTAs under consideration are unlikely to substantially alter production prices, since they only involve a limited share of each partner country's exports. Second, we assume total import demand by product ($P_{jkt}C_{jkt}$) to remain constant in (tax-inclusive) value. We thus neglect the impact of RTAs on income, as well as possible substitutions between imports and domestic products (an alternative assumption is considered below). Under these assumptions, it follows from equation (5) that import demand in value can be expressed as:

$$(10) \quad X_{ijkt} = X_{ijkt_0} \left(\frac{\tau_{ijkt}}{\tau_{ijkt_0}} \right)^{-\sigma} \left(\frac{P_{jkt_0}}{P_{jkt}} \right)^{1-\sigma}.$$

This equation paves the way for counterfactual simulations of the trade impact of RTAs. Let us consider an RTA implemented in year $t_0 + 1$. Assuming all other trade determinants to remain unchanged, preferential tariff cuts in subsequent years are the only source of trade changes, hence:

$$(11) \quad \tilde{X}_{ijkt} = X_{ijkt_0} \left(\frac{\tau_{ijkt}}{\tau_{ijkt_0}} \right)^{-\sigma} \left[\theta_{ijk} \left(\frac{\tau_{ijkt}}{\tau_{ijkt_0}} \right)^{1-\sigma} + \left(1 - \theta_{ijk} \right) \right]^{-1},$$

where, as before, X refers to tax-exclusive import values. *Ceteris paribus*, the RTA implementation shifts trade from X_{ijkt_0} to $\tilde{X}_{ijkt}$. An alternative assumption would be the Armington assumption in its initial form, namely the assumption that products are differentiated by country of origin, with domestic products included in the same bundle (i.e., the same CES function) as other imports. In this case, an approach similar to the one above can be followed, but including domestic consumption. Accordingly, the θs should be defined as value shares in total consumption, not in imports only. This information is not available at the product level, but an upper bound for the corresponding impact can be computed by assuming that this share of imports from the partner in total domestic consumption of the corresponding good is negligible.[15] The counterfactual value of trade is in this case:

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[15] The simulations in Figure 1 refer to the first calculation as the "conservative counterfactual", and to the second as the "upper-bound counterfactual".

(12) $\tilde{\tilde{X}}_{ijkt} = X_{ijkt_0} \left(\frac{\tau_{ijkt}}{\tau_{ijkt_0}} \right)^{-\sigma}$.

Given the assumptions stated above, the conservative counterfactual refers to a case where increases in bilateral exports reflect pure trade diversion, since the total value of imports remains constant. In contrast, trade creation is implicitly assumed to dominate under the upper-bound counterfactual.

Applying these formulas product by product for each exporter-importer pair and aggregating the results allows the consequences of RTAs on agricultural and food exports to be devised. The results presented are based upon our benchmark estimate ($\sigma = 3.71$). Among the 74 agreements studied, only the 57 that entered into force after 1998 are considered in the simulations, due to the need to use trade flows in the year preceding entry into force as a benchmark.

5.2. Simulation results

Not surprisingly, the profile of trade impacts over time and across agreement types parallels those previously found for preferential margins (Figure 3). On average over all the agreements considered, other things being equal, RTAs increase agricultural and food exports by 20% to 28% after five years (as indicated through the conservative and upper-bound counterfactuals, respectively), and by 27% to 39% when they are fully phased in.

The heterogeneity across agreement types, already present when the agreement is being enforced, grows over time. Complete implementation boosts bilateral exporters by 12% to 14% for South-to-North flows, compared to 18% to 31% for flows from the North to the South, and 37% to 52% for flows between Southern countries. Separate results by RTA further illustrate this cross-agreement heterogeneity (available in Appendix). Trade impacts are larger for agreements granting higher preferential margins, in particular when the partner's initial market share is low.

Figure 3. Counterfactual simulations of the impact of RTAs on bilateral exports of agricultural and food products, by type of agreement and years elapsed since entry into force (%)
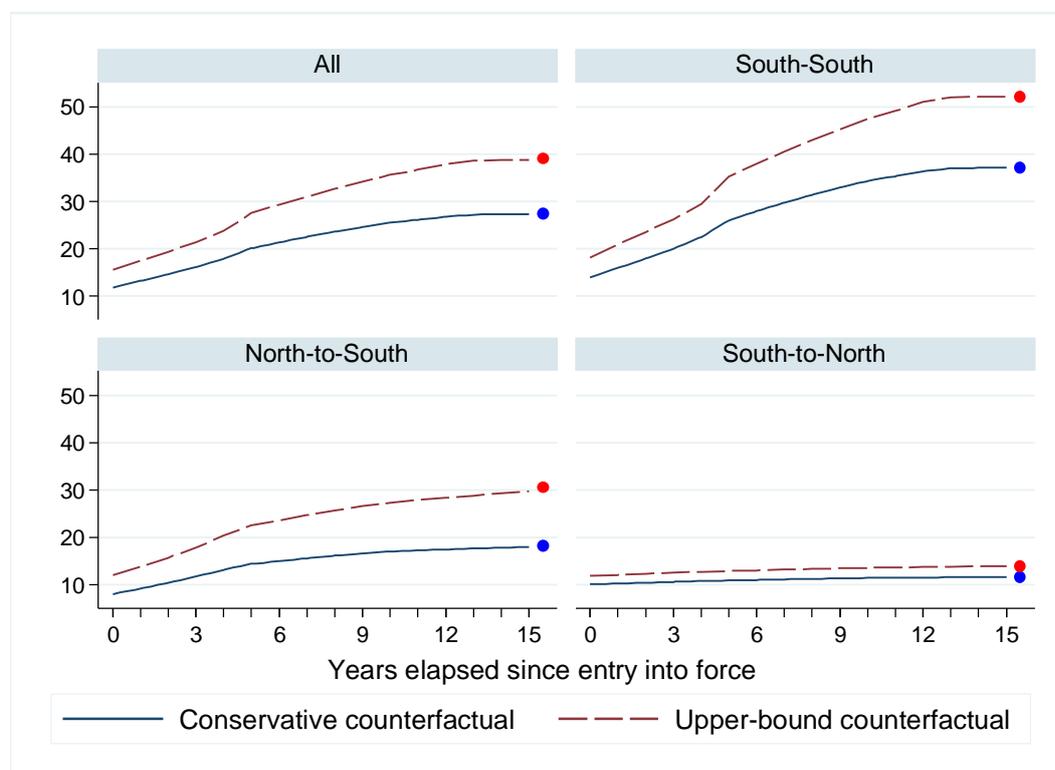
Source: Authors' computations based on equations (11) and (12) and on the benchmark estimate above.

Note: The dots correspond to the full implementation of RTAs, at the end of the phase-in period. The results cover agricultural and food products for 57 agreements entered into force after 1998. TRQ products are included in the simulations, using the inside-quota tariff rate as a measure of protection (excluding them makes little difference, except for South-to-North, where simulated increases are lowered by approximately one quarter –see Appendices for results excluding TRQ products, and for results by individual RTA).

Conclusions

Our analysis demonstrates the increasing importance assumed by RTAs in the international trade of agricultural and food products. Drawing on a database that details tariff concessions on agricultural products in a large number of agreements, it yields a portrait of their incidence in tandem with a progressive implementation which extends over a period exceeding ten years. We find that the preferential margin for agricultural products is approximately 9% in agreements between Southern countries once fully phased-in, while for agreements between

high income countries and others the preferential margins granted by the former are considerably lower (approximately 6% on average) than those granted by the latter (11%).

Estimating the impact of these agreements on trade involves many issues, owing both to the difficulty in correctly specifying the gravity equation that underlies estimates and endogeneity because countries prioritize entering into agreements with partners with whom they have other reasons to desire strong commercial ties. We address these issues by running estimates at the disaggregated level of individual products, only identifying relationships on the basis of trends in trade flows over time, not on cross-sectional comparisons of their intensities. We separately estimate the impact on pre-existing trade flows (the intensive margin) and the impact on the probability of new trade flows arising (the extensive margin). Our estimates confirm the significant impact that preferential agreements have on trade in agricultural products.

We find a positive and significant impact in both cases, at least after the year of entry into force, which it is not possible to conclude based on yearly data. The estimated elasticity of substitution between bilateral trade flows at the product level is found to be 3.7 in our benchmark estimates and the positive impact on the probability of exporting is estimated to be moderate (with a coefficient of one-twentieth). This means that on average a 1 % preferential margin increases pre-existing trade flows by 3.8% and increases the probability to export a given product to a partner country by .05%. When based upon Poisson estimates of the multiplicative model, estimation results appear to be robust to several changes in sample definition. Statistically significant differences are found between types of agreements and between levels of preferential margin, but they remain limited. Heterogeneity appears to be more pronounced across products.

While RTAs have most often been analyzed based on aggregate figures, because of the lack of detailed information, the scope and phasing-in period of tariff preferences varies widely across agreements. Counterfactual simulations show how different the correspondent trade impacts are likely to be. Further research would be useful to document and analyze further these differences and their consequences.

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Appendices

Appendix 1: List of the 74 agreements studied

Argentina-Bolivia (1997); Argentina-Chile (1996); Argentina-Colombia (2005); Argentina-Ecuador (2005); Argentina-Peru (2005); Argentina-Venezuela (2005); Australia-Chile (2009); Australia-Singapore (2003); Australia-Thailand (2005); Australia-USA (2005); Bolivia-Brazil (1997); Bolivia-Mexico (1995); Bolivia-Paraguay (1997); Bolivia-Uruguay (1997); Brazil-Chile (1996); Brazil-Colombia (2005); Brazil-Ecuador (2005); Brazil-Peru (2005); Brazil-Venezuela (2005); Brunei Darussalam-Chile (2006); Canada-Chile (1997); Canada-Costa Rica (2002); Canada-Mexico (1994); Canada-Peru (2009); Chile-China (2006); Chile-Costa Rica (2002); Chile-El Salvador (2002); Chile-Iceland (2004); Chile-Korea (2004); Chile-Mexico (1999); Chile-New Zealand (2006); Chile-Norway (2004); Chile-Paraguay (1996); Chile-Peru (1998); Chile-Singapore (2006); Chile-Switzerland (2004); Chile-Uruguay (1996); Chile-USA (2004); China-New Zealand (2008); Colombia-Paraguay (2005); Colombia-Uruguay (2005); Costa Rica-Mexico (1995); Costa Rica-USA (2006); Dominican Rep.-USA (2006); Ecuador-Paraguay (2005); Ecuador-Uruguay (2005); El Salvador-Mexico (2001); El Salvador-USA (2006); EU27-Chile (2003); EU27-Egypt (2004); EU27-SACU (2000); Guatemala-Mexico (2001); Guatemala-USA (2006); Honduras-Mexico (2001); Honduras-USA (2006); Japan-Mexico (2005); Japan-Thailand (2007); Korea-Singapore (2006); Mexico-Nicaragua (1998); Mexico-Uruguay (2004); Mexico-USA (1994); Morocco-USA (2006); New Zealand-Singapore (2001); New Zealand-Thailand (2005); Nicaragua-USA (2006); Panama-Singapore (2006); Paraguay-Peru (2005); Paraguay-Venezuela (2005); Peru-Thailand (2003); Peru-Uruguay (2005); Peru-USA (2009); Singapore-USA (2004); Turkey-Egypt (2004); Uruguay-Venezuela (2005). The date in parenthesis refers to the year of entry into force of the Agreement.

Appendix 2: Illustrating the statistical reliability of product-level trade flows

The method used requires using control groups' trade flows as a benchmark agains which bilateral exports are gauged. A prerequisite for this approach to be meaningful is that control groups' trade flows be representative. As argued in the text, this is potentially problematic at the product level, in particular for small flows. To illustrate this concern and assess what a reasonable threshold might be to consider product-level trade flows as representative, we start from the premise that a representative flow should exhibit stability over time, since underlying determinants are unlikely to vary radically from one year to the following. The following equation was estimated for control group trade flows:

$$\ln(X_{ijkt}) = \lambda_t + \gamma_g \ln(X_{ijk(t-1)}) + w_{ijkt}$$

Where w_{ijkt} is an error term, λ_t is a set of time dummies, and γ_g is a set of coefficients specific to each category *g*. Trade flows are categorized along *g* depending on the value of X_{ijkt}: lower than USD 10k; 10 to 50k; 50 to 100k; 100 to 200k; 200 to 500k; 500 to 1,000k; more than 1,000k. In Table A.1, this regression is carried out separately for the three types of trade flows used as control to transform the dependent variable: exports from a control group of exporters to a control group of importers; from a control group of exporters to an importer; and from an exporter to a control group of exporters.

In each case, the coefficient of autocorrelations increases with the value of trade flows. In addition, the value of this coefficient is strikingly different for low values than for other flows. This is in particular the case of flows lower than USD 10k, negative in the two cases where there are observations. To a lesser extent, the contrast is also clear for values lower than USD 100k. Given the number of observations in each category (also reported in Table A.1), the threshold of USD 100k seems an acceptable balance between robustness and number of observations.

Table A.1. Estimated autoregressive coefficient of the log of control groups' trade flows, depending on their size

Value of the flow studied (in k USD)	Control group of exporters to control group of importers "I to J"		Control group of exporters to importer "I to j"		Exporter to control group of importers "i to J"	
	Estimate	Nb obs.	Estimate	Nb obs.	Estimate	Nb obs.
]0 ; 10]	na	1	-0.338 ***	59,378	-0.3657 ***	65,061
			0.003		0.0029	
]10 ; 50]	0.513 ***	20	0.192 ***	75,850	0.170 ***	71,831
	(0.147)		(0.002)		(0.002)	
]50 ; 100]	0.804 ***	47	0.398 ***	40,885	0.383 ***	35,488
	(0.022)		(0.002)		(0.002)	
]100 ; 200]	0.846 ***	109	0.491 ***	40,983	0.479 ***	37,087
	(0.024)		(0.002)		(0.002)	
]200 ; 500]	0.890 ***	440	0.570 ***	55,681	0.562 ***	47,761
	(0.009)		(0.001)		(0.001)	
]500 ; 1000]	0.941 ***	1,166	0.627 ***	39,235	0.622 ***	35,389
	(0.004)		(0.001)		(0.001)	
> 1,000	0.987 ***	463,140	0.738 ***	152,911	0.739 ***	172,306
	(0.000)		(0.001)		(0.001)	
Constant	0.147 ***		2.474 ***		2.545 ***	
	(0.004)		(0.008)		(0.008)	
Adj. R-squared	0.98		0.95		0.95	
Observations	351,382		351,382		351,382	

Source: Authors' estimations based on IDB, MacMap-HS6 (ITC and CEPII), BACI (CEPII) and Comtrade (UN Statistics Division) data.

Notes: All estimates are based on equation (1) and include year fixed effects. Only agricultural and food products are covered. The estimation period is 1998-2009. "na", for non available, corresponds to a case for which there is no observation. Robust standard errors are reported in parenthesis. *,**,*** denote statistical significance levels 10%, 5% and 1% respectively. "Nb obs." refers to the number of observations in the estimation sample of our benchmark estimates (larger than the number of observations used here, where the sample is restricted by the use of lagged log exports).

Appendix 3: Robustness checks

Table A.2. Estimation results under alternative assumptions in database construction

	Threshold (in USD) under which control group trade flows are ignored			Multipilicative factor used to filter out instable ratios		
	100,000	200,000 (bench.)	500,000	2	3 (bench.)	4
	(1)	(2)	(3)	(4)	(5)	(6)
FTA's first year	2.19 *	2.22 *	3.69 ***	1.87	2.22 *	2.79 **
	(1.16)	(1.27)	(1.42)	(1.44)	(1.27)	(1.29)
Subsequent years	4.15 ***	3.71 ***	3.27 ***	2.99 ***	3.71 ***	3.70 ***
	(1.09)	(1.16)	(1.08)	(0.96)	(1.16)	(1.04)
Estimation method	PPML	PPML	PPML	PPML	PPML	PPML
Observations	100,398	80,641	54,907	61,369	80,641	88,087
Panel units	13,292	10,703	7,418	9,389	10,703	11,292

Source: Authors' estimates based on IDB, MAcMap-HS6 (ITC and CEPII), BACI (CEPII) and Comtrade (UN Statistics Division) data.

Notes: All estimates include three-way partner-reporter-product fixed effects. Estimates are based on equation (8). The coefficients reported refer to the estimated value of σ, interacted with a dummy indicating whether the FTA is in its first year of implementation (first row) or not (second row). The estimation period is 1998-2009. Robust standard errors are reported in parenthesis, clustered at the panel (reporter-partner-product) level for PPML estimates, at both reporter-year and partner-year level for OLS estimates. Asterisks denote statistical significance levels 10% (*), 5% (**) and 1% (***) respectively.

Table A.3. Estimation results using alternative definition of control groups, or using weights

	Benchmark estimates		Ref. groups restricted by income group		Reference groups restricted by continent		Weighted estimates	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
FTA's first year	2.22 *	0.08	2.13 *	0.15	4.37	-0.28	2.08 ***	0.36
	(1.27)	(0.36)	(1.28)	(0.34)	(2.78)	(0.41)	(0.19)	(0.37)
Subsequent years	3.71 ***	0.94 ***	3.62 ***	0.98 ***	3.69 **	1.28 ***	3.87 ***	1.27 ***
	(1.16)	(0.28)	(1.20)	(0.28)	(1.60)	(0.32)	(0.17)	(0.28)
Estimation method	PPML	OLS	PPML	OLS	PPML	OLS	PPML	OLS
Observations	80,641	63,122	80,050	62,665	57,750	44,912	73,651	63,122
Panel units	10,703	9,462	10,617	9,375	8,555	7,465	9,306	9,462

Source: Authors' estimates based on IDB, MAcMap-HS6 (ITC and CEPII), BACI (CEPII) and Comtrade (UN Statistics Division) data.

Notes: All estimates include three-way partner-reporter-product fixed effects. Estimates are based on equation (8) for PPML estimates and on equation (7) for OLS estimates. The coefficients reported refer to the estimated value of σ, interacted with a dummy indicating whether the FTA is in its first year of implementation (first row) or not (second row). "Low-mean panels" are those partner-reporter-product triplets for which the mean value of non-zero trade flows is lower than USD 100,000 (see text). The estimation period is 1998-2009. Robust standard errors are reported in parenthesis, clustered at the panel (reporter-partner-product) level for PPML estimates, at both reporter-year and partner-year level for OLS estimates. Asterisks denote statistical significance levels 10% (*), 5% (**) and 1% (***) respectively.

"Income group" here refer to the binary taxonomy North v. South described in the text. In estimations (3) and (4), the control group for any country only includes countries from the same income group as the country considered. Continents are grouped into three, namely Eurafrica, Asia-Oceania and America. In estimations (5) and (6), control groups are similarly restricted to countries belonging to the same continent as the one being controlled.

Appendix 4: Counterfactual simulation results, by RTA

Table A.4. Simulated ceteris paribus increase in bilateral exports of agricultural and food products (%)

Country i	Country j	From i to j				From j to i			
		Year 1	Year 5	Full	Full, exc. TRQ	Year 1	Year 5	Full	Full, exc. TRQ
Argentina	Colombia	9.9	24.3	60.1	60.1	30.8	34.0	35.9	35.9
Argentina	Ecuador	5.0	14.9	37.2	37.2	21.9	23.0	23.7	23.7
Argentina	Peru	3.8	8.0	17.3	17.3	17.4	22.7	24.8	24.8
Argentina	Venezuela	5.2	15.4	39.5	39.5	13.3	36.5	49.5	49.5
Australia	Chile	19.6	19.7	19.7	19.7	5.3	5.9	6.0	6.0
Australia	Singapore	1.9	1.9	1.9	1.9	4.8	4.8	4.8	4.8
Australia	Thailand	7.8	14.2	20.4	19.2	5.5	5.5	5.5	5.5
Brazil	Colombia	10.8	23.0	40.9	40.9	13.7	23.1	24.3	24.3
Brazil	Ecuador	5.9	19.0	50.0	50.0	52.1	54.0	54.5	54.5
Brazil	Peru	4.4	14.0	34.9	34.9	22.2	27.9	30.6	30.6
Brazil	Venezuela	8.1	22.8	57.8	57.8	21.8	29.9	30.4	30.4
Canada	Costa Rica	4.5	10.1	15.1	15.1	3.4	3.4	3.4	3.4
Canada	Peru	22.9	23.2	23.4	23.4	8.5	8.6	8.6	8.6
Chile	China	8.7	26.3	46.9	46.9	13.2	14.2	14.3	14.3
Chile	Costa Rica	24.3	28.4	34.8	34.4	16.9	16.9	16.9	16.9
Chile	El Salvador	25.8	29.7	35.3	35.3	23.8	23.8	23.8	23.8
Chile	Mexico	50.3	56.3	60.2	56.2	14.5	14.5	14.5	14.5
Colombia	Paraguay	9.2	31.1	89.3	89.3	13.3	20.3	36.7	36.7
EU27	Chile	9.5	11.1	11.3	11.5	7.8	18.4	19.4	19.9
EU27	Egypt	1.5	1.5	1.5	0.9	1.7	1.7	1.7	3.4
EU27	SACU	5.5	7.8	16.7	16.7	12.2	14.2	16.5	18.4
Iceland	Chile	23.9	23.9	23.9	23.9	4.1	4.1	4.1	4.1
Japan	Mexico	4.6	12.3	12.6	14.2	12.6	13.7	13.9	7.3
Japan	Thailand	14.4	48.2	62.9	66.8	5.5	7.4	9.1	9.7
Korea	Chile	19.8	19.9	20.0	20.0	12.0	37.8	76.2	76.5
Mexico	El Salvador	5.2	13.3	21.6	21.6	12.8	15.2	17.5	17.5
Mexico	Guatemala	2.6	7.0	12.0	12.0	16.8	30.2	39.9	39.9
Mexico	Honduras	4.2	11.2	19.3	19.3	8.4	21.0	31.7	31.7
Mexico	Uruguay	26.0	26.0	26.0	26.0	102.5	102.5	102.5	154.3

Table A.4 (cont'd)

Country i	Country j	From i to j				From j to i			
		Year 1	Year 5	Full	Full, exc. TRQ	Year 1	Year 5	Full	Full, exc. TRQ
New Zealand	Chile	12.7	14.6	17.3	17.3	5.4	5.4	5.4	5.4
New Zealand	China	7.0	20.1	24.5	28.9	4.2	7.9	7.9	7.9
New Zealand	Thailand	9.6	19.1	27.2	35.3	10.9	10.9	10.9	10.9
Norway	Chile	0.0	0.0	0.0	0.0	0.1	0.1	0.1	0.1
Paraguay	Ecuador	3.0	9.9	28.6	28.6	14.3	28.5	40.7	40.7
Peru	Paraguay	11.7	40.8	94.4	94.4	3.4	10.7	17.9	17.9
Singapore	Chile	22.8	22.8	22.8	22.8	0.3	0.3	0.3	0.3
Singapore	Korea	18.0	192.1	219.4	219.4	113.9	113.9	113.9	113.9
Singapore	New Zealand	7.5	7.5	7.5	7.5	0.4	0.4	0.4	0.4
Singapore	Panama	10.7	27.2	29.6	29.6	0.0	0.0	0.0	0.0
Switzerland	Chile	2.8	2.8	2.8	2.8	0.8	0.8	0.8	0.8
Thailand	Peru	0.0	0.1	0.1	0.1	71.9	106.7	121.0	121.0
Turkey	Egypt	2.3	2.3	2.3	3.8	0.1	0.1	0.1	0.4
Uruguay	Colombia	18.8	25.2	39.0	39.0	12.6	27.0	65.3	65.3
Uruguay	Ecuador	0.1	0.2	0.3	0.3	11.7	13.9	16.6	16.6
Uruguay	Peru	5.9	17.0	40.1	40.1	1.4	1.4	1.4	1.4
Uruguay	Venezuela	3.3	12.1	29.4	29.4	8.4	27.9	60.1	60.1
USA	Australia	6.8	6.8	6.8	6.8	10.8	10.9	15.1	5.3
USA	Chile	6.8	8.7	11.2	11.8	2.3	3.3	5.6	3.7
USA	Costa Rica	5.6	8.7	14.9	12.6	10.4	10.4	10.7	8.3
USA	Dominican Rep.	4.6	7.8	13.2	11.8	29.9	30.7	32.8	21.1
USA	El Salvador	6.2	9.0	15.9	14.3	25.1	25.1	25.4	3.7
USA	Guatemala	6.4	9.7	14.5	11.7	20.1	20.5	21.9	13.3
USA	Honduras	6.5	9.3	15.0	16.7	16.6	17.0	18.3	16.2
USA	Morocco	24.6	47.8	54.6	52.2	8.8	9.4	10.8	6.8
USA	Nicaragua	6.2	10.1	18.2	12.3	29.5	30.2	32.7	10.0
USA	Peru	19.7	22.5	26.9	22.7	15.1	15.2	15.4	13.1
USA	Singapore	1.4	1.4	1.4	1.4	9.7	14.0	18.6	3.0

Note: Results based on conservative counterfactual simulations, using equation (11). In each row, the first named country ("i") is the one with higher PPP GDP per capita in 2001. Rows ranked by alphabetical order of the first named country. "Year 1" refers to the year following the entry into force of the agreement, "Year 5" to the fifth year after entry into force, "Full" to the full implementation of RTAs, once the phase-in period is over. Except in "Full, exc. TRQ" columns, TRQ products are included in the simulations, using the inside-quota tariff rate as a measure of protection. Only agreements entered into force after 1998 are covered in this table.