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REVISITING THE EFFECTS OF REGIONAL TRADING AGREEMENTS ON TRADE FLOWS WITH PROPER SPECIFICATION OF THE GRAVITY MODEL

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Abstract
This paper uses a gravity model to assess ex-post regional trade agreements. The model includes 130 countries and is estimated with panel data over the period 1962-1996. The introduction of the correct number of dummy variables allows for identification of Vinerian trade creation and trade diversion effects, while the estimation method takes into account a potential correlation between some of the explanatory variables and the bilateral specific effects introduced in the model, as well as a potential selection bias. In contrast to previous estimates in cross-section, results show that regional agreements have generated a significant increase in trade between members, often at the expense of the rest of the world.

Résumé
Ce papier présente une évaluation ex-post des effets des accords régionaux à l'aide d'un modèle de gravité estimé en panel, sur un échantillon incluant 130 pays sur la période 1962-1996. L'introduction dans le modèle de 3 ensembles de variables muettes permet d'identifier les effets de création et détournement de trafic (selon la terminologie de Viner), et la méthode d'estimation utilisée corrige l'endogénéité de certaines variables explicatives ainsi qu'un potentiel biais de sélection. Contrairement aux résultats des estimations en transversal, il apparaît que les accords régionaux ont engendré une augmentation significative du commerce entre leurs membres, souvent au détriment du reste du monde.

JEL Classification: F11; F15; C23.

Keywords: Regional trade agreements, Gravity equation, Trade creation, Trade diversion, Panel Data.

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1. INTRODUCTION

After a long period of neglect from the late sixties (Poyhonen, 1963, Tinbergen, 1962, Linnemann, 1966) to the late eighties, the gravity trade model has acquired a second youth. First, new theoretical foundations have been proposed both with the advent of trade theories based on increasing returns to scale, in imperfectly competitive markets and firm-level product differentiation (Helpman and Krugman, 1985, Bergstrand 1985, 1989, Baier and Bergstrand, 2001 or Evenett and Keller, 2002) and, within a perfect competition setting, with product differentiation at the national level (Deardorff, 1998 or Anderson and Van Wincoop, 2003). Second, the gravity model has been used extensively to study trade patterns, as for example in the case of the drastic changes following the demise of central planning. Most recently, in the estimation of models of geography and trade, the gravity model is, once again, holding center stage (Hummels, 1999, Redding and Venables, 2001, Limao and Venables, 2001). In fact, the gravity model has also become a favored tool to assess the ex-post trade effects of a currency union (Glick and Rose, 2002 or Rose and Van Wincoop, 2001), or the trade creating (TC) and trade diverting (TD) effects associated with preferential trading arrangements (Frankel, 1997 or Soloaga and Winters, 2001). However, for reasons elaborated in this paper, previous estimates of TD and TC are likely to be unreliable.

Along with this renewal in interest, questions have been raised about the proper formulation of the model (choice of variables) as well as about proper econometric techniques, especially when the usual cross-country formulation is amended to include a temporal dimension. Indeed, the discussion about the proper econometric specification of the gravity model has shown that the conventional cross-section formulation without the inclusion of country specific effects is misspecified and so introduces a bias in the assessment of the effects of regional trading agreements (RTAs) on bilateral trade (e.g., Matyas, 1997, Soloaga and Winters, 2001). However, it turns out that this panel specification, with three specific effects (exporter, importer and time effects) is only a restricted version of a more general model which allows for country pair heterogeneity (e.g., Cheng and Wall, 1999 or Egger and Pfaffermayr, 2000).

In contrast to the traditional cross-section gravity model which includes time invariant trade impediment measures (e.g. distance, common language dummies, border, historical and cultural links as in most studies, see Frankel, 1997), this general proposed specification is more adequate since it accounts for any time invariant (unobserved) bilateral effect. Hence, all factors that influence bilateral trade which were partially captured by regional dummies are now controlled for.

In this paper, I apply this more general panel specification on a recent gravity model specification derived by Baier and Bergstrand (2002) with the addition of: (i) a barrier-to-trade function similar to Limao and Venables (2001) instead of the traditional distance variable and common border dummy, and: (ii) three dummy variables for each RTA considered (intra-trade, imports and exports dummies) to allow for a correct identification of Vinerian trade effects. I show that the predictions of the effects of RTAs in terms of trade creation and trade diversion are very different whether one uses a cross-section or a panel specification with random bilateral effects (fixed effects eliminating agreements that are time invariant). In this setting, one has to check for the potential correlation of some explanatory variables with the country-pair unobservable effects. I show that the use of the instrumental variable method proposed by Hausman and Taylor (1981) is necessary to avoid estimation bias. Moreover, the selection bias that can appear in an unbalanced sample is tested and corrected for by the inclusion of a selection rule in the model's estimation (e.g. Nijman and Verbeek, 1992 or Guillotin and Sevestre, 1994).

Section 2 presents the canonical gravity model (with the barrier-to-trade function) and the modified cross-section version used for ex-post evaluations of RTAs (with the three dummies mentioned above that have to be included for each RTA). Section 3 specifies the alternative panel model with the characteristics proposed above (bilateral random effects, and correction for the endogenous explanatory variables and for the selection bias). Finally, section 4 presents the average effects of each RTA over 1962-1996 and the evolution of these effects over the same period, comparing cross-section and panel estimates. To anticipate the main conclusion, it turns out that the panel estimates yield more convincing estimations in the average effects (as RTA dummies do not capture anymore unobservable bilateral trade patterns) as in the evolution which suggest that, globally, around the date of implementation, RTAs generated larger increases in trade among members than predicted with cross-section estimates. Insofar as the gravity model is the accepted model for estimating the efficiency effects of RTAs, attention should be paid to the proper specification of the model. Section 5 concludes.

2. THE GRAVITY MODEL AS AN EX-POST METHOD TO ASSESS REGIONAL AGREEMENTS

2.1 The canonical gravity model

Following the recent exposition in Baier and Bergstrand (2002), I use a generalized version of the "standard" gravity model derived from a framework where firms maximize profits and consumers maximize utility according to Dixit-Stiglitz preferences. As shown by Baier and Bergstrand (2002), if the representative profit-maximizing firms in country j set product prices delivered to market i according to equation (2), one obtains the following equilibrium trade flow for each goods-producing firm in country j to market i:

$$M_{ij} = A\frac{Y_j}{p_j}Y_i\left[\frac{p_j\theta_{ij}}{P_i}\right]^{1-\sigma}\left[s_j\left(1+t_i\right)\left(1+t_{ij}\right)^{-\sigma}\right] \tag{1}$$

M_{ij} : c.i.f value of the aggregate merchandise trade flow imported by country i from exporter j;

A: parameter that captures technology and taste parameters which are assumed identical across countries and are absorbed in a constant term in the estimated equation;

$Y_{i(j)}$: gross domestic product of country i (j);

p_j : exporter (country j) price level of it representative good. The price level of this good in country i (c.i.f price) is given by:

$$p_{ij} = p_j\theta_{ij} \tag{2}$$

with θ_{ij} a barrier-to-trade function between i and j to be developed below;

P_i: output-weighted measure of the remoteness (in terms of trade costs) of country i

$$P_i = \left[\sum_{k=1}^{N} n_k\left[p_k\theta_{ik}\left(1+t_{ik}\right)\right]^{1-\sigma}\right] \tag{3}$$

with n_j being the number of varieties of goods produced in j;

t_{ij}: ad valorem tariff rate by country i on the good produced in j (t_{ii}=0 assumed);

s_j: real share of goods output in national product in country j;

t_i: share of tariff revenue relative to income;

Equation (1) is the currently accepted theoretical foundation for the gravity equation in the presence of transportation costs and tariffs, cf. Feenstra (2003, chap.5). Assuming $t_i = 0$ (in most countries, tariff revenue is a trivial share of GDP), and following Anderson and Van Wincoop (2003) who have shown the implicit solution for p_j and P_i, yields:

$$M_{ij} = \frac{A}{Y_w} s_j Y_i Y_j \theta_{ij}^{1-\sigma} \left(1+t_{ij}\right)^{-\sigma} \left[P_i^* P_j^*\right]^{\sigma-1} \tag{4}$$

where Y_w is world output of goods and P_i* and P_j* can be interpreted as "multilateral price resistance terms" (see e.g. Anderson and Van Wincoop, 2003).

Equation (4) is remarkably close to the gravity model in empirical literature. It suggests that the proper specification should include:

(1) the logarithm of the product of the GDPs of countries i (Y_i) and j (Y_j);

(2) per capita GDP or population of the exporting country, N_j, as a proxy for the capital-endowment ratio (which determines the endogenous share of goods in national output i.e. s_j)[2].

(3) a proxy for the term θ_{ij}. It is obvious that it is crucial to get the best handle possible on what constitutes the "barriers-to-trade" function, which is usually proxied either by distance, D_{ij}, between trading partners (and the presence of a common border and a common language as for instance in Baier and Bergstrand, 2002 or Anderson and Van Wincoop, 2003), or sometimes by the c.i.f/f.o.b price ratio[3]. Because recent studies have shown that these variables are not the only determinants of barriers to trade, I model the barrier-to-trade function, between countries i and j, as follows[4]:

$$\theta_{ij} = (D_{ij})^{\delta_1} (IN_i)^{\delta_2} (IN_j)^{\delta_3} \left[e^{\delta_4 L_{ij} + \delta_5 E_i + \delta_6 E_i}\right] \tag{5}$$

with (expected signs on coefficients in parentheses):

[2] Populations (or alternatively per capita incomes) now have a straightforward interpretation in the gravity equation: higher populations (for given incomes) reduce the capital-labor endowment ratios of the exporter countries, tending to reduce the capital-intensive industry's share of national output. Then, merchandise trade flow (the dependent variable) should fall, relative to national output, as goods are capital intensive (e.g. Baier and Bergstrand, 2002).

[3] Baier and Bergstrand (2001) use the c.i.f/f.o.b ratio to model transport costs, but their study only deals with OECD countries which have better data. For a discussion about the problems associated with the use of c.i.f/f.o.b data see Hummels (1999) and Limao and Venables (2001).

D_{ij}: distance between countries i and j (δ_1>0);

L_{ij} = 1 if i and j share a common border, otherwise 0 (δ_4<0);

$E_{i(j)}$= 1 if the country i (j) is landlocked; otherwise 0 (δ_5>0, δ_6>0);

$IN_{i(j)}$: level of infrastructure of the country i (j), computed as an average of the density of road, railway and the number of telephone lines per capita (δ_2<0, δ_3<0).

(4) the product of the multilateral resistance term for the country pair. As discussed in Anderson and Van Wincoop (2003), proper estimation of P_i* and P_j* requires a nonlinear estimation technique[5]. We use the following multilateral resistance or "remoteness" proxies[6]:

$$R_i = \left[\sum_{k=1, k \neq i}^{N} Y_k \left(D_{ik} \right)^{1-\sigma} \right]^{\frac{1}{1-\sigma}} \tag{6a}$$

$$R_j = \left[\sum_{k=1, k \neq j}^{N} Y_k \left(D_{kj} \right)^{1-\sigma} \right]^{\frac{1}{1-\sigma}} \tag{6b}$$

where (6a) and (6b) are estimated using a central elasticity value σ=4 (results in the paper are robust to 2<σ<6).

Taking into account the modifications discussed above, after substitution of (5) in (4), the estimated reduced-form boils down to:

$$\begin{aligned} \ln M_{ij} &= \beta_0 + \beta_1 \ln Y_i + \beta_2 \ln Y_j + \beta_3 \ln N_j + \beta_4 \ln R_i + \beta_5 \ln R_j + \beta_6 \ln D_{ij} \\ &+ \beta_7 \ln L_{ij} + \beta_8 \ln E_i + \beta_9 \ln E_j + \beta_{10} \ln IN_i + \beta_{11} \ln IN_j + \omega_{ij} \end{aligned} \tag{7}$$

where $\frac{A}{Y_w}$ is absorbed in the constant term, and with the expected signs:

β_1>0, β_2>0, β_3<0, β_4>0, β_5>0, β_6=(1-σ)δ_1<0, β_7=(1-σ)δ_4>0, β_8=(1-σ)δ_5<0, β_9=(1-σ)δ_6<0, β_{10}=(1-σ)δ_2>0 and β_{11}=(1-σ)δ_3>0.

[4] e.g. Limao and Venables (2001).

[5] Measures of multilateral resistance (cf. Anderson and Van Wincoop 2003) or fixed effects (cf. Rose and Van Wincoop 2001 and Feenstra 2003) must be included to avoid an omitted variables bias. Anderson and van Wincoop (2003) and Feenstra (2003) show that both approaches yield consistent estimates in a gravity equation.

[6] Polak (1996) shows that if one doesn't use a measure of the average distance between a country and its main partners as well as absolute distance in assessing the effects of RTAs, one will underestimate trade between faraway countries and thus bias the RTA coefficient. As shown in the recent derivations of the gravity model, one should include the average distance of the importing country i from its main partners.

2.2 The gravity model for ex-post assessment of regional trade agreements

Initially used by Aitken (1973) as an ex-post assessment for the EEC, the gravity model seems well defined for this issue for two reasons. Firstly, arguably, the model represents a relevant counterfactual to isolate the effects of an RTA. If the sample of countries is appropriately selected, the gravity equation then suggests a "normal" level of bilateral trade for the sample. Then, dummy variables can be used to capture the "atypical" levels resulting from an RTA. Secondly, thanks to the correct introduction of dummy variables in the model, one can isolate trade creation (TC) and trade diversion (TD) effects of an RTA.

In a Vinerian world following a RTA, TC and TD will be reflected in trade flows as follows :(i) under pure TC, intra-regional trade increases and imports from the rest of the world (ROW) remain unchanged; (ii) under pure TD, the increase in intra-regional trade is entirely offset by a corresponding decrease in imports from the ROW; (iii) if there is both TC and TD, intra-regional trade increases more than imports from the ROW decrease. Because of second-best considerations, identification of TD and TC does not allow inference about the welfare consequences of a RTA for its members. Finally, for non-members, one should include a measure of the change in volume of exports from members to non-members (an increase signifying an improvement in welfare for non-members)[7].

Therefore, the correct ex-post assessment (e.g. Endoh (1999) and Soloaga and Winters (2001)) of a RTA on the volume of trade should include the following dummy variables (associated coefficients in parentheses)[8]:

(i) D_I (α_I) =1 if both partners belong to the same RTA [zero otherwise] (capturing intra-bloc trade);

(ii) D_M (α_M) =1 if importing country i belongs to the RTA and exporting country j, to the ROW [zero otherwise] (capturing bloc imports from the ROW);

(iii) D_X (α_X) =1 if exporting country j belongs to the RTA and importing country i to the ROW [zero otherwise] (capturing bloc exports to the ROW).

[7] see Winters (1997). Note that, if looking at the exports to the ROW is more important than imports from the ROW in terms of non-member welfare, the critical welfare variable is the terms of trade.

[8] Most authors (e.g. Bayoumi and Eichengreen, 1997, Frankel, 1997, Krueger, 1999) have not included enough dummy variables to distinguish between exports and imports, so they fail to isolate TD and TC effects.

Suppose that $\alpha_I > 0$, which corresponds to more intra-bloc trade than predicted by the reference ($\alpha_I < 0$ corresponding to a RTA between complementary economies) and which can be in substitution to domestic production or to exports from the ROW. Hence, to conclude on whether this corresponds to TC or TD, one needs to examine the signs of the coefficients α_M and α_X. Then, $\alpha_I > 0$ along with a lower propensity to import from the ROW ($\alpha_M < 0$) indicates TD, and if the increase in intra-regional trade is entirely offset by a decrease in regional imports from the ROW, we have pure TD. If intra-regional trade increases more than imports from the ROW decrease, there is both TC and TD. And with $\alpha_I > 0$ and $\alpha_M \geq 0$, there is pure TC. Finally, comparing α_I and α_X can lead to inferences about welfare for non-members. For example, ($\alpha_I > 0, \alpha_X < 0$) would indicate a dominant "export diversion" and hence a decrease in welfare for non-members.

To summarize, following a RTA, [$\alpha_I > 0$ and $\alpha_M \geq 0$ ($\alpha_X \geq 0$)] indicates pure TC in terms of imports (exports) and [$\alpha_I > 0$ and $\alpha_M < 0$ ($\alpha_X < 0$)], indicates TD in terms of imports (exports).

3. DATA AND ESTIMATION

The model is estimated with data for 130 countries over the period 1962-1996. Trade data are from UN COMTRADE (total bilateral imports in current dollars). Data sources for the explanatory variables along with data transformations are presented in appendix A.1. Once the missing values are taken out[9], the sample covers 130 countries (a list of the countries in the sample is presented in appendix A.2). There are thus 240 691 observations for 14 387 pairs of countries.

3.1 Panel specification

It has been observed repeatedly (see Polak, 1996, Matyas, 1997, Bayoumi and Eichengreen, 1997) that regional dummy variables in cross-country estimates capture everything specific to the importing or exporting countries not captured by the variables included in the equation that influence the level of trade (e.g. historical,

[9] Countries which do not declare their imports from a partner or which do not import from this partner are identified in the same way, i.e. with a missing value. Hence, our data are not censored at zero. The actual number of observations (240 691) represents around 50% of potential number. This implies a potential selection bias which is tested (and corrected for) in section 3.3.

cultural, ethnic, political or geographical factors)[10] which is troublesome since the dummy variables should really isolate TD and TC effects. Not taking into account the countries' heterogeneity or of the pair of countries in bilateral trade relations may introduce a bias. By contrast, a panel data method enables one to identify the specific effects of the pair of countries and to isolate them. The usual correction introduces three specific effects: exporter, importer and time effects (e.g. Matyas, 1997 and Soloaga and Winters, 2001). But the model with three specific effects is only a restricted version of the more general model which allows for country-pair heterogeneity adopted here (e.g. Cheng and Wall, 1999 and Egger and Pfaffermayr, 2000). Then a bilateral term, μ_{ij}, specific to each pair of countries and common to each year (and different according to the direction of trade: $\mu_{ij} \neq \mu_{ji}$), is included in the previous model (7) specified in panel:

$$\begin{aligned}\ln M_{ijt} = {} & \alpha_0 + \alpha_t + \beta_1 \ln Y_{it} + \beta_2 \ln Y_{jt} + \beta_3 \ln N_{jt} + \beta_4 \ln R_{it} + \beta_5 \ln R_{jt} + \beta_6 \ln D_{ij} \\ & + \beta_7 \ln L_{ij} + \beta_8 \ln E_i + \beta_9 \ln E_j + \beta_{10} \ln IN_{it} + \beta_{11} \ln IN_{jt} + \beta_{12} \ln RER_{ijt} + \mu_{ij} + \nu_{ijt}\end{aligned} \tag{8}$$

α_0 : effect common to all years and pairs of countries (constant);

α_t : effect specific to year t but common to all the pairs of countries to capture common shocks (e.g. changes in oil prices);

μ_{ij} : effect specific to each pair of countries and common to all the years.

Note the introduction of the bilateral real exchange rate (RER_{ijt}) in equation (8). In a model with panel data that span a long time period (here 35 years), it is essential to capture the evolution of competitiveness (e.g. Soloaga and Winters, 2001 and Bayoumi and Eichengreen, 1997). An increase in the bilateral real exchange rate reflecting a depreciation of the importing country's currency against that of the exporting country, one would expect $\beta_{12}<0$.

Following the specification check, the three dummy variables discussed above are introduced in the model to detect TD and TC for a selection of RTAs.

[10] If these factors are also correlated with gravity variables (GDP, populations, distance), estimations which do not include them will have an endogeneity bias, because the omitted variables are correlated with the level of bilateral trade and with the explanatory variables (see below).

3.2 Econometric method

The RTAs considered in the paper are: EU, ANDEAN, NAFTA, CACM, MERCOSUR, ASEAN and LAIA (see appendix A.2 for definition and members). To assess the total effect of these RTAs, we consider the trade between RTAs members (and with the ROW) even before the implementation of the agreements, in order to look for break points around the important dates of the agreements (and notably before and after the implementation date). Hence, a fixed effects model is inadequate for the estimation of the effects of NAFTA, CACM, MERCOSUR, ASEAN and LAIA as there is no change in their members[11] (the Within transformation eliminates time-invariant variables). Then, bilateral effects are modeled as random variables. In the absence of correlation between the explanatory variables and the specific bilateral effects, the GLS estimation provides consistent estimates of the coefficients. However, variables like GDPs, infrastructure or even RTA intra-trade may be correlated with bilateral specific effects[12].

The usual way to deal with this issue is to consider instrumental variables estimation such as that proposed by Hausman and Taylor (henceforth HT) (1981), though here it is adapted to the case of an unbalanced sample according to the method proposed by Guillotin and Sevestre (1994). Let X (Z) the matrix of explanatory variables variant (invariant) over time and suppose that among the variables X and Z, there exist: (i) $X_{ijt:}$ k_1 (k_2) exogenous (endogenous) variables, denoted X_1 (X_2);

(ii) $Z_{ij:}$ g_1 (g_2) exogenous (endogenous) variables denoted Z_1 (Z_2).

If the condition $k_1 \geq g_2$ is satisfied, then the equation is identified and (8) can be estimated using $[QX_1, QX_2, PX_1, Z_1]$[13] as instruments (see Breusch, Mizon and Schmidt, 1989). As the resulting estimator is consistent but not efficient[14], HT suggest using this first round of estimates to compute the variance of the specific effect (σ_μ^2), and the variance of the error term (σ_v^2)[15]. The instrumental variable estimator is then applied to the following transformed equation:

[11] In contrast, for the UE and the ANDEAN, as they have been "enlarged" within the time span 1962-1996, the fixed effect estimator can be used (see later, column 1 table 1).

[12] The Hausman test (1978) allows us to control for the presence of correlation between explanatory variables and specific bilateral effects.

[13] Q is the matrix that computes the deviations from individual means. P is the matrix that computes the observation across time for each individual (pair of countries).

[14] as it does not correct for heteroskedasticity and serial correlation due to the presence of random bilateral specific effects.

[15] with σ_μ^2 and σ_v^2 corrected for the bias of heteroskedasticity specific to the unbalanced sample.

$$M_{ijt} - (1\text{-}\Omega_{ij})\, M_{ij.} = [X_{ijt} - (1\text{-}\Omega_{ij})X_{ij}]\ \psi + \Omega_{ij}\ Z_{ij}\phi + \Omega_{ij}\ \mu_{ij} + [\,\nu_{ijt} - (1\text{-}\Omega_{ij})\ \nu_{ij.}] \qquad (9)$$

With[16] $\Omega_{ij} = (\sigma_\nu^2 / T_{ij}\sigma_\mu^2 + \sigma_\nu^2)^{1/2}$.

A Hausman test of over-identification, based on the comparison of the HT estimator and the Within estimator, must be carried out. If the null hypothesis cannot be rejected, the instruments are legitimate (in the sense of no bias due to a correlation between specific bilateral effects and the explanatory variables), and the HT estimator is the most efficient estimator (e.g. Baltagi, 2001). Note that canonical correlations are also a useful device for comparing different sets of instruments[17]. In fact, one should use instruments for which the geometric average of the canonical correlations with the regressors are maximized (e.g. Baltagi, 2001, Mairesse *et al.,* 1999 or Hall *et al.* 1996).

3.3 Endogeneity of explanatory variables and sample selection bias

I check first for endogeneity among explanatory variables.[18] Results are reported in table 1. Column 1 in table 1 reports estimates from the Within equation which treats the bilateral specific effects as fixed, thereby giving unbiased parameter estimates for time-varying variables. All these coefficients are significant at the 1% level and have the expected sign. The fit is good (R^2=0.87) and the specific bilateral and time effects introduced in the model are strongly significant (as showed by the Fisher tests).

Next come the results from the error-component model (GLS) which differ markedly from the Within estimation results. The Hausman test, based on differences between Within and GLS estimators, reveals a χ^2_7 = 944.39, which is highly significant (1%). Hence, this test rejects the null hypothesis according to which there would be no correlation between the bilateral specific effects and the explanatory variables. The GLS estimator is thus biased, and the use of the HT method is therefore justified.

For sensitivity analysis, five regressions are estimated with the HT estimator. The over-identification test indicates for each regression if the instruments are legitimate

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[16]The average value of Ω_{ij} will be systematically presented in the tables of results.
[17] I thank an anonymous referee for suggesting using canonical correlations.
[18] Because the dataset covers a long time span, some series may contain a unit root and thus the estimates in the table 1 may be spurious. So, a Levine and Lin (1993) unit root test has been applied to the series for GDP, population and bilateral import. This test rejects, for all series, the null of a unit root.

or if an additional source of correlation between specific effects and explanatory variables exists (in the case of a significant test statistic).

Insert Table 1 here

The first estimation, labeled HT I in table 1, considers only the GDP variables (Y_{it} and Y_{jt}) as endogenous. The results point out that these variables are actually correlated with the specific effects: the Hausman test, which compares HT to GLS, confirms that the instrumentation has improved the model[19] (the hypothesis of exogeneity of GDP variables is rejected). However, the over-identification test rejects the hypothesis according to which there would be no more correlation between explanatory variables and bilateral effects ($\chi^2_5 = 224.22$). Hence, only a part of the initial bias has been corrected.

A second source of correlation can come from the population variables. Equation HT II takes these two variables (and the GDP variables) as endogenous. The corresponding tests for this equation lead us to conclude that once again, the model has been improved but the difference with the Within estimation is still significant.

A third source of endogeneity could come from infrastructure variables[20]. Their instrumentation, in addition to those for income and population, improves the model and the over-identification test indicates that the hypothesis of legitimacy of the instruments used cannot be rejected. As the identification condition is verified, the HT estimator is convergent and more efficient than the Within estimator.

However, as noted[21] by Baier and Bergstrand (2002, p.5), "the FTA dummy variable may be endogenous by being correlated with unobservable (omitted) variables that are correlated also with the decision to trade". Actually, if cultural, political or historical ties between countries increase their propensity to form a trade agreement as well as their bilateral volume of trade, then there would still be a bias in the coefficient for

[19] Guillotin and Sevestre (1994) recommend comparing the HT estimator to the GLS estimator with a Hausman test. If the null H_0 is rejected, we can conclude that the instrumented model gives better estimations then the GLS model (without any instrumentation). Thus, instrumented variables are actually endogenous.

[20] I have also tested for the correlation of distance and remoteness variables with bilateral effects in the equation HT III. However, this equation does not improve the model HT II.

[21] I thank an anonymous referee for this suggestion.

intra-RTA trade. HT V instruments the seven intra-RTA trade dummies[22]. There is no change in the coefficients for the traditional gravity variables, but the coefficients for intra-RTA trade are significantly higher (compared to HT IV) for four RTAs. The over-identification test remains good and the geometric average of the canonical correlations (0.70 for HT IV and 0.75 for HT V) gives an idea of the gains of asymptotic efficiency as one moves from HT IV to HT V[23].

Concerning traditional gravity variables, coefficients have the expected sign and are significant at a 1% level (except for E_i). Import volume of i from j increases with GDP and coefficients are close to unity as suggested by the theory. The exporting country population variable has the expected negative sign. The elasticity of bilateral trade to distance is close to unity (-0.91)[24] and the coefficients for the remoteness variables are significantly positive. The volume of trade increases with the level of infrastructure of each country, as in Limao and Venables (2001). Sharing a land border allows countries to trade 2.9 times more than expected from the gravity equation ($=e^{1.07}$). Likewise, imports from a country without direct access to the sea are 36% lower. Finally, a real depreciation of i with respect to j lowers i's imports from j. These results are sensible and overall, very comforting.

However, a last potential estimation bias deserves scrutiny: the unbalanced sample could be subject to a non-ignorable selection rule[25]. In this case, the selection bias can be tested and corrected by the inclusion of the selection rule in the model estimation. I use a method proposed by Nijman and Verbeek (1992), which approximates the Heckman correction term by adding variables that reflect the individual's patterns in terms of presence in the sample to the model (see Guillotin and Sevestre, 1994). So HT IV is estimated again including the following additional variables: (i) PRES: number of years of presence of the couple ij's in the sample;(ii) DD: dummy that takes

[22] One can use a linear projection to instrument for endogenous binary variables as mentioned in Wooldridge (2000, p.84).

[23] According to the Barghava and al. Durbin Watson test (1982), modified to the unbalanced panel, the HT IV and HT V residuals are not autocorrelated AR(1): there is no systematic difference between observed and predicted trade flows. Hence, the HT estimator is efficient and the over-identification test is appropriate.

[24] According to equation (8), $\beta_6=(1-\sigma).\delta_1$. The elasticity of transport costs with respect to distance is usually estimated in the range $0.2<\delta_1<0.4$ (e.g. Limao and Venables, 2001, Hummels, 1999). Combined with an elasticity of substitution between goods of about 4 (Rose and Van Wincoop, 2001, use an estimate of σ=5; Obstfeld and Rogoff, 2001, suggest a consensus estimate of σ between 4 and 6), the implied distance coefficient would be in the range $-0.8<\beta_6<1.2$ which is almost identical to the estimates of β_6 in table 1.

[25] i.e. that the probability of a pair of countries being included in the sample is not independent of model error, and in particular to the unobserved bilateral effects.

the value 1 if ij is observed during the entire period, 0 otherwise;(iii) PA_t: dummy that takes the value 1 if ij was present in t-1 (PA_0=0).

Results from this estimation are reported in the first column of table 2 and are compared to equation HT V (table 1). The conclusions of the previous estimations are not modified, even if the coefficients associated to the variables PRES and PA_t are significantly different from zero. Hence, *ceteris paribus*, pairs of countries which have at least two years of consecutive available data (and *a fortiori* if they are present over several years) have more bilateral trade than pairs of countries with interruption in their data. These three variables will be systematically introduced in future regressions, in order to avoid the selection bias in the coefficients of regional dummies.

4. APPLICATION TO THE ASSESSMENT OF THE EFFECTS OF REGIONAL TRADE AGREEMENTS

I proceed in two steps: first I comment the average effects of each RTA over the entire sample period (1962-1996) reported in table 2, then I decompose these average effects to look for the evolutions over the period (notably around the important dates).

4.1 Average effects over the period 1962-1996

Table 2 reports the coefficients for dummy variables for two sets of regressions, one in cross-section (corresponding to most uses of the gravity model for ex-post assessments of RTAs), yielding 35 separate regressions (one for each year), the other with the panel specification of section 3.

Insert Table 2 here

Results are quite different depending on the estimation method. For instance, on average over 1962-1996, intra-EU trade is 104% (=$100(e^{0.713}-1)$) above what is predicted by the panel gravity model, whereas it is 21% below the expected level according to the cross-section gravity model. The latter negative result is obtained in other cross-section studies, i.e. Frankel (1997) or Soloaga and Winters (2001). Likewise, in the panel estimation, the ANDEAN and the ASEAN blocs present a trade between members about respectively 1.2 and 2.4 times above the reference prediction, associated with a propensity to import from the ROW inferior by 62% and 38%. By contrast, in the cross-section estimations, the positive intra-trade for these two RTAs is associated, on average, with a positive propensity to import from the ROW.

Since relevant inferences about TD and TC require inspection of the evolution of these coefficients over time and around the period when RTAs are implemented, I break down regional dummy variables into two-year periods with these variables replacing the global regional dummies.[26]

4.2 Evolution of the effects during the RTA's existence

Because the results are self-explanatory from inspection of figure 1 to 3, I comment the EU results and give only an overall interpretation for NAFTA and MERCOSUR along with a summary for other RTAs[27]. Concerning the changes in dummy variables commented in this section, note that all the non significant coefficients are graphed as zero and when two consecutive coefficients are significant, the change between these two coefficients is always tested statistically significant in the panel estimations.

First of all, the cross-section analysis (figure 1a) displays a negative trend in intra-EU trade until 1980 before it turns positive with the propensity to export to the ROW declining over the period, suggesting exports TD but no evidence of import-TD. This result is similar to what Soloaga and Winters (2001) obtained using the same estimation method.

Insert Figures 1 here

By contrast, panel estimates (figure 1b) suggest three rather distinct periods in terms of TC and TD. From 1967 to 1973, intra-trade presents no clear tendency, as for the trade with the ROW. However, following the first (and second) enlargement, the model predicts a significant positive trend in intra-trade (α_I increases and turns significantly positive in 1978, the pattern continuing with the deep integration following the EC-92 program). In parallel, there is first a stagnation of imports of

[26] Each RTA dummy is multiplied by 18 time dummies of two years (except the last one that capture only the year 1996). Coefficient estimates for explanatory variables are identical to the previous ones under this procedure because for each agreement, the addition of the new variables introduced is equal to the former aggregate dummy variable.
[27] The evolutions of the estimated coefficients are presented in appendix A.3 for cross-section results and for panel ones for the other RTAs.

members from the ROW until 1985 and then a negative trend (α_M became negative in 1990). Hence, the model suggests that, if the first enlargement of the EU (from six to nine members in 1974) resulted in a pure TC, the second enlargement (with Spain and Portugal in 1986 and subsequent deep integration) presents sign of significant TD, in terms of imports and exports. Note however that deep integration in the form of reduced technical barriers to trade, even if discriminatory, cannot give rise to welfare reduction for RTA members. These results are quite different from Bayoumi and Eichengreen (1997), and arguably more sensible, since many studies have raised concerns about overall TD effects because of the effects of the common agricultural policy.

Insert Figures 2 and 3 here

For the EU, the sample period did not allow the inclusion of a pre-RTA period. This is not the case for the other RTAs. Comparing the results from both estimation methods is even more striking in the cases of MERCOSUR and NAFTA. Here, the cross-section estimates show largely unexplainable volatility throughout the time-period whereas the panel estimates capture much more clearly the expected effects of a RTA around the time of announcement or of implementation: an increase in intra-trade and a decrease in imports from the ROW. The difference in patterns is particularly striking for NAFTA, which reveals largely insignificant dummies until the first trade policy reforms in Mexico, and the announcement of NAFTA negotiations. As to MERCOSUR, panel estimates capture both the increase in intra-trade and the diversion of import from the ROW captured in the more disaggregated analysis in Yeats (1998). At the same time, there is some evidence of an increase of exports for NAFTA and MERCOSUR to the ROW (which probably reflects reduction in trade barriers to non-partners at the same time as they were forming the RTA). Clearly, the panel estimates reveal a more plausible pattern than the cross-section estimates, and yield plausible patterns, which is not the case for the cross-section estimates.

This pattern of import (and sometimes export) TD was also found for other RTAs reported in appendix A.3. For example, in the case of the ANDEAN agreement, the model finds import-TD over the period 1969-1981, over the period 1962-1974 for the CACM, and over the period 1968-1980 for the LAIA. Concurrently, over the same period, an export-TD is observed for the ANDEAN, whereas there is some evidence of an increase of the propensity to export towards the ROW for CACM. ASEAN and LAIA (after 1980) are the only examples of pure TC over the period.

5. CONCLUSIONS

This paper has paid particular attention to the specification and the estimation of the gravity model to correct for biases present in previous studies. The panel estimation with bilateral specific random effects was revealed to be statistically justified after correcting for the endogeneity of the income, size, infrastructure and intra-RTA trade variables. Moreover, dummies were introduced to take into account the selection rule of the sample. Arguably, these modifications lead to a better formulation of the counterfactual against which one assesses the trade performance of RTAs.

Comparison of panel estimates with the more usual cross-section estimates revealed a far more plausible pattern of trade effects associated with RTAs as evidenced by the examination of three well-studied RTAs: EU, MERCOSUR and NAFTA. In general, the findings of this study, covering seven RTAs, show that most of these RTAs resulted in an increase in intra-regional trade beyond levels predicted by the gravity model, often coupled with a reduction in imports from the rest of the world, and at times coupled with a reduction in exports to the rest of the world, suggesting evidence of trade diversion.

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Table 1. Results of the estimates of the gravity equation on panel data.

Variables	M_{ijt}						
	Within	**GLS**	**HT I a)**	**HT II b)**	**HT III c)**	**HT IV d)**	**HT V e)**
ln Y_{it}	1.09**	0.82**	1.00**	1.01**	1.01**	1.08**	1.09**
	(58.1)	*(114.2)*	*(96.7)*	*(89.8)*	*(92.0)*	*(99.0)*	*(87.2)*
ln Y_{jt}	1.15**	1.06**	1.15**	1.11**	1.16**	1.15**	1.15**
	(62.6)	*(116.2)*	*(104.9)*	*(98.7)*	*(101.9)*	*(91.4)*	*(87.7)*
ln N_{jt}	-0.60**	-0.17	-0.15**	-0.57**	-0.45**	-0.59**	-0.60**
	(-17.2)	*(-15.4)*	*(-12.5)*	*(-21.6)*	*(-17.4)*	*(-22.8)*	*(-21.7)*
ln R_{it}	0.13**	0.01	0.11**	0.12**	0.12**	0.12**	0.13**
	(12.5)	*(0.9)*	*(12.0)*	*(12.4)*	*(12.8)*	*(13.2)*	*(13.5)*
ln R_{jt}	0.25**	0.39**	0.08**	0.21**	0.20**	0.25**	0.21**
	(11.7)	*(30.15)*	*(4.41)*	*(11.2)*	*(10.3)*	*(13.6)*	*(11.2)*
ln D_{ij}	-	-1.00**	-1.01**	-0.89**	-0.27**	-0.95**	-0.92**
		(-48.6)	*(-44.7)*	*(-37.4)*	*(-2.7)*	*(-41.7)*	*(-29.6)*
L_{ij}	-	1.18**	1.18**	1.45**	2.56**	1.08**	1.07**
		(11.3)	*(10.8)*	*(12.6)*	*(12.3)*	*(12.4)*	*(9.8)*
E_i	-	0.03	-0.10	0.02	0.01	-0.03	0.04
		(0.6)	*(-1.9)*	*(0.4)*	*(0.2)*	*(-0.5)*	*(0.6)*
E_j	-	-0.19**	-0.40**	-0.48**	-0.42**	-0.43**	-0.44**
		(-4.3)	*(-8.3)*	*(-9.4)*	*(-7.5)*	*(-8.6)*	*(-6.7)*
ln IN_{it}	0.04**	0.07**	0.02**	0.02**	0.03**	0.04**	0.04**
	(6.5)	*(11.8)*	*(3.8)*	*(3.5)*	*(4.9)*	*(5.3)*	*(6.0)*
ln IN_{jt}	0.03**	0.01**	0.01**	0.02**	0.01**	0.03**	0.02**
	(5.0)	*(3.2)*	*(2.9)*	*(3.7)*	*(3.1)*	*(6.6)*	*(5.3)*
ln RER_{ijt}	-0.006**	-0.006**	-0.005**	-0.006**	-0.006**	-0.006**	-0.006**
	(-4.4)	*(-4.7)*	*(-3.8)*	*(-4.7)*	*(-4.5)*	*(-4.8)*	*(-4.7)*
Number of obs (NT)	240 691	240 691	240 691	240 691	240 691	240 691	240 691
Number of bilateral (N)	14 387	14 387	14 387	14 387	14 387	14 387	14 387
R^2 f)	0.87	0.66	0.65	0.64	0.59	0.65	0.69
Ω_{ij} (mean)	-	0.82	0.82	0.83	0.84	0.82	0.84
Bilateral fixed effect	35.50** *F(14386,226256)*	-	-	-	-	-	-
Time fixed effect	71.72** *F(34,226256)*	165.79** *F(34,240623)*	125.98** *F(34,240623)*	115.67** *F(34,240623)*	150.50** *F(34,240623)*	156.41** *F(34,240623)*	171.73** *F(34,240623)*
Hausman test W vs. GLS g) *chi-2(Kw)*	-	944.39** *chi-2 (14)*	-	-	-	-	-
Hausman test HT vs. GLS h) *chi-2(K)*	-	-	1115.26** *chi-2(33)*	1268.52** *chi-2(33)*	1594.23** *chi-2(33)*	1479.64** *chi-2(33)*	1514.37** *chi-2(33)*
Test of over-identification i) *chi-2 (k_1-g_2)*	-	-	224.22** *chi-2(12)*	40.34** *chi-2(11)*	211.21** *chi-2(8)*	3.22 *chi-2(9)*	0.87 *chi-2(2)*
Canonical correlation j)	-	-	0.81	0.62	0.76	0.70	0.75

** and * significant at the 1% and 5% levels respectively (t-student is presented under the correspondent coefficient).
The time dummy variables and the constant are not reported in order to save space.
[a] HT I : endogenous variables = lnY_{it} and lnY_{jt}.
[b] HT II : endogenous variables = lnY_{it}, lnY_{jt}, and lnN_{jt}.
[c] HT III : endogenous variables =lnY_{it}, lnY_{jt}, lnN_{jt}, lnR_{it}, lnR_{jt} and lnD_{ij}.
[d] HT IV : endogenous variables = lnY_{it}, lnY_{jt}, lnN_{jt}, $lnIN_{it}$ and $lnIN_{jt}$.
[e] HT V : endogenous variables = lnY_{it}, lnY_{jt}, lnN_{jt}, $lnIN_{it}$, $lnIN_{jt}$, and the 7 RTA in dummies.
[f] Calculated, for GLS and HT, from [1-Sum of Square Residuals] / [Total Sum of Squares] on the transformed model. Note that the impact of random specific effects are not in the R^2 but are part of residuals.
[g] This test is applied to the differences between the Within and GLS estimators, without taking into account the coefficients of time effects. If we take them into account, the result is: chi-2(48)= 1373.65**
[h] Hausman test applied to the differences between GLS and HT estimators, without time effects.
[i] Hausman test applied to the differences between Within and HT estimators, without time effects.
[j] Geometric average of the canonical correlation coefficients.

Table 1 end Results of the estimates of the gravity equation on panel data.

Variables	M_{ijt}						
	Within	**GLS**	**HT I a)**	**HT II b)**	**HT III c)**	**HT IV d)**	**HT V e)**
EU_{intra}	0.56**	0.91**	0.25**	0.28**	0.27**	0.30**	0.74**
	(24.4)	*(18.9)*	*(5.1)*	*(5.7)*	*(5.5)*	*(6.1)*	*(15.0)*
$EU_{imports}$	0.27**	0.07**	0.15*	0.15	0.13	0.24*	0.22**
	(13.8)	*(3.8)*	*(2.2)*	*(1.0)*	*(1.8)*	*(2.1)*	*(11.3)*
$EU_{exports}$	0.26**	0.41**	0.36**	0.34**	0.35**	0.34**	0.32**
	(14.0)	*(17.9)*	*(15.7)*	*(14.8)*	*(15.2)*	*(14.6)*	*(13.8)*
$ANDEAN_{intra}$	0.59**	0.98**	0.10	0.15	0.14	0.75**	0.60**
	(6.2)	*(4.8)*	*(0.8)*	*(0.4)*	*(1.2)*	*(6.1)*	*(4.7)*
$ANDEAN_{imports}$	-0.93**	-0.82**	-0.53**	-0.53**	-0.58**	-0.84**	-0.95**
	(21.6)	*(-14.6)*	*(-14.4)*	*(-14.2)*	*(-14.9)*	*(-22.7)*	*(-24.8)*
$ANDEAN_{exports}$	-0.95**	-0.69**	-0.60**	-0.51**	-0.62**	-0.83**	-0.92**
	(19.8)	*(-11.3)*	*(-14.8)*	*(-12.4)*	*(-14.4)*	*(-20.2)*	*(-21.9)*
$CACM_{intra}$	-	0.58**	0.77**	0.62**	0.99**	0.69**	0.78**
		(6.7)	*(7.1)*	*(6.5)*	*(8.4)*	*(6.9)*	*(5.2)*
$CACM_{imports}$	-	-0.42**	-0.33**	-0.30**	-0.49**	-0.62**	-0.77**
		(-5.0)	*(-3.7)*	*(-3.2)*	*(-4.8)*	*(-6.9)*	*(-6.3)*
$CACM_{exports}$	-	0.11	0.10	-0.17	-0.18	-0.05	-0.08
		(1.3)	*(1.1)*	*(-1.8)*	*(-1.7)*	*(-0.5)*	*(-0.6)*
$LAFTA_{intra}$ **a)**	-	0.58**	0.45**	0.48**	0.82**	0.52**	0.56**
		(8.7)	*(7.0)*	*(6.9)*	*(7.7)*	*(7.3)*	*(5.5)*
$LAFTA_{imports}$	-	-0.57**	-1.47**	-1.54**	-1.89**	-1.51**	-1.49**
		(-9.1)	*(-11.9)*	*(-12.0)*	*(-12.7)*	*(-12.2)*	*(-8.7)*
$LAFTA_{exports}$	-	-0.13	-0.56**	-0.35*	-0.77	-0.36**	-0.31
		(-1.84)	*(-4.04)*	*(-2.4)*	*(-4.6)*	*(-2.6)*	*(-1.6)*
$MERCOSUR_{intra}$	-	-0.38	-0.50	-0.32	-0.21	-0.41	-0.35
		(0.8)	*(1.0)*	*(-0.6)*	*(0.4)*	*(0.8)*	*(0.5)*
$MERCOSUR_{imports}$	-	-1.23**	-1.38**	-1.43**	-1.86**	-1.03**	-1.09**
		(-14.3)	*(-14.8)*	*(-14.8)*	*(-15.5)*	*(-15.3)*	*(-8.6)*
$MERCOSUR_{exports}$	-	-0.17	-0.31	0.01	-0.49**	-0.09	-0.03
		(-1.9)	*(-3.4)*	*(0.2)*	*(-4.1)*	*(-1.0)*	*(-0.2)*
$NAFTA_{intra}$	-	0.08	0.01	0.78	0.60	0.42	0.38
		(0.11)	*(0.0)*	*(1.0)*	*(0.8)*	*(0.6)*	*(0.4)*
$NAFTA_{imports}$	-	-0.26**	-0.30**	-0.43**	-0.58**	-0.39**	-0.51**
		(-3.0)	*(-3.0)*	*(-4.1)*	*(-5.0)*	*(-3.8)*	*(-3.8)*
$NAFTA_{exports}$	-	0.43**	0.55**	0.73**	0.23	0.11	-0.01
		(3.8)	*(3.6)*	*(6.1)*	*(1.8)*	*(1.1)*	*(0.1)*
$ASEAN_{intra}$	-	2.02**	1.74**	1.27**	0.71**	0.64**	0.81**
		(5.3)	*(4.5)*	*(5.6)*	*(6.1)*	*(5.4)*	*(4.5)*
$ASEAN_{imports}$	-	-0.32**	-0.25**	-0.21**	-0.12	-0.52**	-0.49**
		(-4.5)	*(-3.3)*	*(-2.7)*	*(-1.2)*	*(-2.9)*	*(-4.7)*
$ASEAN_{exports}$	-	1.05**	1.15**	1.14**	0.63**	0.75**	0.69**
		(12.6)	*(10.4)*	*(13.2)*	*(5.9)*	*(10.1)*	*(8.8)*

** and * significant at the 1% and 5% levels respectively (t-student is presented under the correspondent coefficient).

[a)] As all the members of ANDEAN and MERCOSUR belong also to LAIA, we isolate the evolution of trade of the two former RTA in computing the dummies for LAIA as follows (i.e. Soloaga and Winters, 2001):

$LAIA_{intra}$ =LAIA intra- $ANDEAN_{intra}$- $MERCOSUR_{intra}$;

$LAIA_{imports}$= LAIA imports- $ANDEAN_{imports}$- $MERCOSUR_{imports}$;

$LAIA_{exports}$= LAIA exports- $ANDEAN_{exports}$- $MERCOSUR_{exports}$.

Table 2. Comparison of the results with regional dummies (1962-1996).

Variables	M_{ijt}				
	Panel HT V		Cross-section		
	Coeff.	t	Average Coeff.	max	min
ln Y_{it}	1.086**	*78.1*	0.799	*0.95***	*0.68***
ln Y_{jt}	1.146**	*84.2*	1.003	*1.26***	*0.76***
ln N_{jt}	-0.665**	*-23.7*	-0.083	*0.03*	*-0.22***
ln R_{it}	0.131**	*14.0*	0.154	*0.27***	*0.07*
ln R_{jt}	0.193**	*10.2*	0.099	*0.57***	*0.00*
ln D_{ij}	-0.977**	*-22.9*	-1.035	*-0.54***	*-1.29***
L_{ij}	1.088**	*5.2*	0.846	*1.48***	*0.45***
E_i	-0.134	*-1.5*	-0.118	*0.15*	*-0.52***
E_j	-0.332*	*-3.4*	-0.348	*-0.02*	*-1.10***
ln IN_{it}	0.036**	*5.4*	0.201	*0.47***	*-0.09*
ln IN_{jt}	0.025**	*5.8*	0.405	*0.63***	*0.19**
ln RER_{ijt}	-0.005**	*-4.1*	-	-	-
PRES	0.046**	*14.2*	-	-	-
DD	-0.086	*-0.5*	-	-	-
PA_t	0.478**	*49.1*	-	-	-
EU $_{intra}$	0.713**	*14.6*	-0.239	*0.44**	*-1.02***
EU $_{imports}$	0.203**	*10.5*	0.896	*1.09***	*0.55***
EU $_{exports}$	0.294**	*12.6*	1.017	*1.62***	*0.26**
ANDEAN $_{intra}$	0.598**	*4.8*	1.351	*2.31***	*0.06*
ANDEAN $_{imports}$	-0.966**	*-25.0*	0.425	*1.18***	*-0.86***
ANDEAN $_{exports}$	-0.916**	*-21.4*	-0.490	*0.23*	*-1.01***
CACM $_{intra}$	0.792**	*2.9*	2.270	*2.79***	*1.60***
CACM $_{imports}$	-0.797**	*-4.7*	-0.474	*0.14*	*-0.87***
CACM $_{exports}$	-0.209	*-1.2*	-0.152	*0.37**	*-0.91***
LAIA $_{intra}$	0.547**	*2.7*	1.225	*1.60***	*-0.07*
LAIA $_{imports}$	-1.580**	*-6.6*	-1.051	*-0.46**	*-1.87***
LAIA $_{exports}$	-0.423	*-1.6*	-0.576	*0.66***	*-2.22***
MERCOSUR $_{intra}$	-0.903	*-0.9*	-0.516	*1.18***	*-1.68***
MERCOSUR $_{imports}$	-1.094**	*-9.1*	0.011	*0.62***	*-0.54***
MERCOSUR $_{exports}$	-0.179	*-1.0*	0.136	*1.16***	*-0.92***
NAFTA $_{intra}$	0.647	*0.48*	0.770	*2.15***	*-0.50**
NAFTA $_{imports}$	-0.497**	*-2.7*	0.319	*0.68***	*0.12*
NAFTA $_{exports}$	0.058	*1.1*	0.072	*0.63***	*-0.65***
ASEAN $_{intra}$	0.881**	*5.7*	1.495	*2.40***	*1.11***
ASEAN $_{imports}$	-0.476**	*-3.5*	0.343	*0.89***	*0.01*
ASEAN $_{exports}$	0.755**	*9.0*	0.469	*1.20***	*-0.25*
Number of obs (NT)	240 691		6 877	*8 472*	*6 012*
Number of bilateral (N)	14 387		-	-	-
R^2	0.69		0.65	*0.73*	*0.60*
Ω_{ij} (mean)	0.91		-	-	-

** and * significant at the 1% and 5% levels respectively (t-student next to correspondent coefficient).

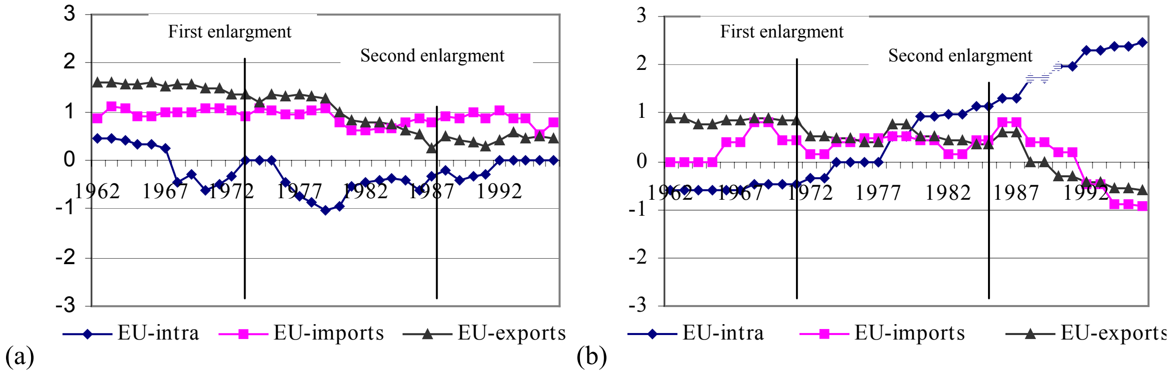

Fig. 1. Evolution of EU dummies over 1962-1996 (α_I, α_M and α_X):
(a) in cross-section and (b) in panel.

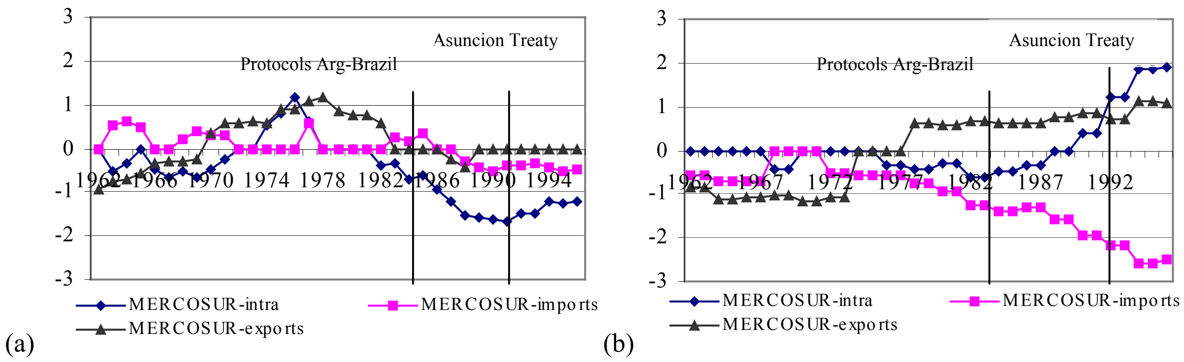

Fig. 2. Evolution of MERCOSUR dummies over 1962-1996 (α_I, α_M and α_X):
(a) in cross-section and (b) in panel.

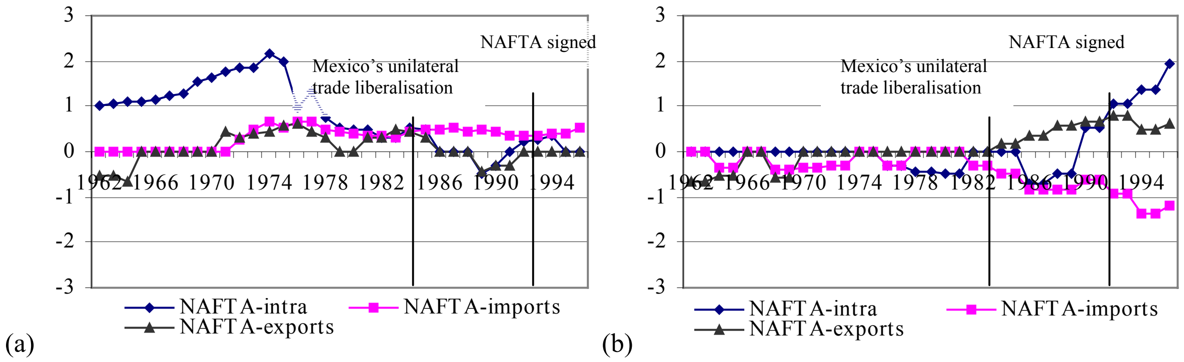

Fig. 3. Evolution of NAFTA dummies over 1962-1996 (α_I, α_M and α_X):
(a) in cross-section and (b) in panel.

Appendices to "Revisiting the Effects of Regional Trading Agreements on trade flows with Proper Specification of the Gravity Model"

A.1 : Sources and definitions of data.

$\mathbf{M_{ijt}}$ **:** COMTRADE, aggregate total bilateral imports of country i from country j (current US $).

$\mathbf{Y_{i(j)t}}$ **:** CD-ROM WDI, World Bank 1999, GDP of country i at time t (constant US $ 1995).

$\mathbf{N_{it}}$ **:** CD-ROM WDI, World Bank 1999, total population of country i at time t.

$\mathbf{D_{ij}}$ **:** Data for distance are extracted from the software developed by the company CVN. The distance is orthodromic, i.e. taking into account the sphericity of Earth, and measured in kilometers between the main economic city of the country i and that of country j.

$\mathbf{L_{ij}}$ **:** Dummy equal to one if the countries i and j share a common land border, 0 otherwise.

$\mathbf{E_{i(j)}}$ **:** Dummy equal to one if the country i (j) does not have a direct access to the sea, 0 otherwise.

$\mathbf{IN_{i(j)t}}$ **:** This index is built using 4 variables from the database constructed by Canning (1996): the density of roads, of paved roads, of railways, and the number of telephone lines per capita of country i (j) at time t, each variable being normalized to have a mean equal to one. An arithmetic average is then calculated over the four variables, for each country and each year (the computation is similar to Limao and Venables, 2001).

$\mathbf{R_{i(j)t}}$ **:** Multilateral trade resistance or "remoteness" of country i (j), computed according to the equation (6a) and (6b) with σ=4.

$\mathbf{RER_{ijt}}$ **:** We extract from the IFS data set the nominal exchange rate for each country against US $ ($NER_{i/\$}$, country i's currency value of 1 US$), and the consumption price index for country i (CPI_i), for each year from 1962 to 1996. If the CPI is not available for a country, we consider the GDP deflator of the country. The bilateral real exchange rate (RER) is computed as following:

$RER_{i/j} = (CPI_j) / (CPI_i) . (NER_{i/\$} / NER_{j/\$})$.

For each pair of countries, we specify the RER such as its mean over the period is zero.

A.2 Countries in the sample and definition of the RTAs studied.

OECD	Sub-Saharan Africa	Latin America and the Caribbean	Asia and the Pacific	Others
Australia	Angola	Argentina[e,i]	Bangladesh	*Albania*
Austria[a]	South Africa*	Bahamas	Brunei	*Armenia*
Belgium +	Burundi	Barbados	Bhutan	*Azerbaijan*
Luxembourg[a]	Benin	Belize	China	Bulgaria
Canada[f]	Burkina Faso	Bolivia[b,e]	Fiji	*Belarus*
Germany[a]	Central African Rep.	Brazil[e,i]	Hong Kong	Czech Rep.
Denmark[a]	Ivory Coast	Chile[b,e]	Indonesia[c]	Algeria
Spain[a]	Cameroon	Colombia[b,e]	India	Saudi Arabia
Finland[a]	Congo	Costa Rica[d]	*Cambodia*	Egypt
France[a]	*Comoros*	Dominican Rep.	*Lao PDR*	Estonia
U. K.[a]	Cape Verde	Dominica	Macao	*Georgia*
Ireland[a]	Djibouti	Ecuador[b,e]	*Mongolia*	Greece[a]
Iceland	Ethiopia + Eritrea	Grenada	Malaysia[c]	Bosnia and
Italy[a]	Gabon	Guatemala[d]	*Nepal*	Herzegovina
Japan	Ghana	Guyana	Pakistan	Hungary
Korea, Rep.	Guinea	Honduras[d]	Philippines[c]	Iran
United States[f]	Guinea-Bissau	Haiti	Papua New Guinea	Israel
Netherlands[a]	Gambia	Jamaica	Singapore[c]	Jordan
Norway	Equatorial Guinea	Mexico[e,f]	Salomon Islands	*Kazakstan*
New Zealand	Kenya	Nicaragua[d]	Thailand[c]	*Kyrgyz Rep.*
Portugal[a]	Madagascar	Panama	*Vietnam*	Kuwait
Sweden[a]	Mali	Peru[b,e]	*Western Samoa*	Lithuania
Switzerland +	Mozambique	Paraguay[e,i]	Sri Lanka	Latvia
Liechtenstein	Mauritania	El Salvador[d]	*Tonga*	*Macedonia*
	Mauritius	Suriname	*Kiribati*	Morocco
	Malawi	Trinidad and Tobago	*Vanuatu*	Malta
	Niger	Uruguay[e,i]		Oman
	Nigeria	St. Vincent and		Poland
	Rwanda	The Grenadines		Romania
	Sudan	Venezuela[b,e]		*Russian Federation*
	Senegal	*St. Lucia*		Slovenia
	Sierra Leone	*Antigua and*		Slovak Rep.
	Sao Tomé and Principe	*Barbuda*		Syrian Rep.
	Seychelles	*St. Kitts and*		*Tajikistan*
	Somalia	*Nevis*		*Turkmenistan*
	Chad			Tunisia
	Togo			Turkey
	Tanzania			*Ukraine*
	Uganda			*Uzbekistan*
	Zaire			
	Zambia			
	Zimbabwe			

Countries written in italic are not available as reporter countries in COMTRADE (only as partners).

* South Africa includes bilateral trade of the group of countries: South Africa + Lesotho + Botswana + Namibia + Swaziland.

a) EU member (European Union, 1957) except U.K., Denmark, Ireland (1973), Greece (1981), Spain, Portugal (1986), Austria, Finland and Sweden (1995);
b) ANDEAN member (1969), except Chile (1969-1976), Venezuela (1973), Peru (1969-1992);
c) ASEAN member (Association of Southeast Asian Nations, 1967);
d) CACM member (Central American Common Market, 1960);
e) LAIA member (Latin American Integration Association, 1980 + former LAFTA, 1960);
f) NAFTA member (North American Free Trade Agreement, 1992);
i) MERCOSUR member (Mercado comùn del Sur, 1991).

A.3. Evolution of the RTA dummies estimated in panel and in cross-section over 1962-1996.