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Revisiting Regional Trading Agreements

with Proper Specification of the Gravity Model

Céline CARRERE*

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* CERDI : Centre d'études et de recherches sur le développement international,
65 boulevard François Mitterrand,
63 000 Clermont Ferrand – France.
Tel : (33) 473 177 400. Fax : (33) 473 177 428.
E.mail : C.Carrere@cerdi.u-clermont1.fr.

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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 in panel over the period 1962-96. 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 the explanatory variables and the bilateral specific effects introduced in the model, as well as potential selection bias. In contrast with previous estimates, results show that over the period 1962-1996, regional agreements have generated a significant increase in trade between members, often at the expense of the rest of the world.

JEL Classification: F11; F15; C23.

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

Résumé

Ce papier utilise un modèle de gravité pour évaluer ex-post des accords commerciaux régionaux. Le modèle est estimé en panel, sur 130 pays et sur la période 1962-96. L'introduction du nombre correct de variables muettes permet d'identifier les effets de création et de détournement de trafic vinérien, selon une méthode d'estimation qui prend en compte (i) la corrélation potentielle entre certaines variables explicatives et les effets spécifiques bilatéraux introduits dans le modèle, (ii) un biais de sélection potentiel. Contrairement aux estimations des études précédentes, les résultats mettent en évidence que sur la période 1962-1996, les accords régionaux considérés dans ce papier ont engendré une augmentation significative du commerce entre les pays membres, souvent au détriment du reste du monde.

JEL Classification: F11; F15; C23.

Mots-clé : Accords Commerciaux Régionaux, Equation de gravité, création de trafic, détournement de trafic, données de panel.

1. Introduction

After a long period of neglect following its introduction in the late sixties (Poyhonen, 1963, Tinbergen, 1962, Linnemann, 1966, Aitken, 1973) since the late eighties, the gravity trade model has acquired a second youth. First, it discovered new theoretical foundations both with the advent of the trade theories based on monopolistic competition and firm-level product differentiation which predict that the intensity of trade should be inversely related to GDP across trading partners (Krugman and Helpman, 1985, Feenstra, Markusen and Rose, 1998) and within a perfect competition setting with product differentiation at the national level (Deardorff, 1998). Second, the gravity model is 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, 2001, Redding and Venables, 2001, Limao and Venables, 2001). In fact, the gravity model has also become a favored tool to assess ex-post the trade creating and trade diverting effects associated with preferential trading arrangements (Frankel, 1997, Soloaga and Winters, 2001).

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 agreements on bilateral trade (e.g., Matyas, 1997 and Soloaga and Winters, 2001). However, it turns out that these specifications, 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 and 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 Frankel, 1997), the more general proposed specification is more adequate since it accounts for any (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 specification and show that the predictions of the effects of regional trade agreements (RTA) in terms of trade creation and trade diversion are very different according to 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, the potential correlation of some explanatory variables with the country-pair effects has to be analyzed. I show that the use of the instrumental 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 estimation (as in Guillotin and Sevestre, 1994).

Section 2 presents the canonical gravity model with the modified cross-section version used for ex-post evaluations of regional agreements (with the three dummies mentioned above that have to be included for each RTA according to trade theories). Section 3 specifies the alternative panel model with the characteristics proposed above. Finally, section 4 compares cross-section and panel estimates. To anticipate the main conclusion, it turns out that the panel estimates yield more convincing estimates which also suggest that, globally, RTAs generated larger increases in trade among members than predicted with cross-section estimates. Section 5 concludes.

2. The gravity model as an ex-post method to assess regional agreements

2.1 The standard gravity model

Although several models yield a gravity-type equation, in a framework that emphasizes aggregate trade, it is convenient to derive the gravity equation from a perfect-competition H-O type model under the assumption of complete specialization at the country level, along with product differentiation at the country level. Assume then maximization of a CES utility function (where σ is the common elasticity of substitution between any pair of countries' products, $\sigma>0$). As shown in appendix A.1, this yields the standard "generalized" gravity equation:

$$(1) \qquad M_{ji}=\frac{Y_i Y_j}{Y^W}\left[\frac{\dfrac{(p_i)^{-\sigma}}{(\overline{P}_i)^{1-\sigma}}}{\sum_h \gamma_h \left(\dfrac{p_i}{\overline{P}_h}\right)^{1-\sigma}}\right]=\frac{Y_i Y_j}{Y^W}\theta_{ij}^{-\sigma} e_{ij}^{-\sigma}\left[\frac{\dfrac{(p_j)^{-\sigma}}{(\overline{P}_i)^{1-\sigma}}}{\sum_h \gamma_h \left(\dfrac{p_i}{\overline{P}_h}\right)^{1-\sigma}}\right] \qquad \forall i,j,h=1..n$$

where γ_h is the share of country h in world income, $\overline{P}_i$ is the CES price aggregator in importer country i and p_i is the price in the country of destination i facing consumers. Assume now that the relationship between the price in the country of origin j, p_j, and the country of destination i, p_i is given by :

$$(2) \qquad p_i=p_j\, e_{ij}\, \theta_{ij}$$

In (2), e_{ij} represents the nominal bilateral exchange rate (defined so that an increase in its value corresponds to a depreciation of i's currency with respect to j's currency) and θ_{ij} is a barrier-to-trade function between i and j to be developed below.

To obtain an estimable model from equation (1), three issues need to be considered. First, distance must be measured correctly. Select units of goods so that each country's product price, p_j, is normalized to unity (and e_{ij}=1). Then, as shown by Deardorff (1998), $\overline{P_i}$ (given by equation A3 in appendix A.1) becomes a "CES index of country i's barriers-to-trade factors" as an importer. Hence, if θ_{ij} is proxied by distance between i and j, $DIST_{ij}$, we have to introduce, in addition to the variable of absolute distance between i and j, a variable of average distance of the importing country i from its main partners, ($\overline{DIST_i}$) to take account of "the relative distance of i from suppliers" as suggested by the theoretical models of Anderson (1979) and Deardorff (1998)[1]. Omitting this variable would have important consequences, in particular in assessing the effects of RTAs[2].

Second, it is crucial to get the best handle possible on what constitutes the 'barriers-to-trade' function, θ_{ij} which are usually proxied either by distance, $DIST_{ij}$ between trading partners (and the presence of a common border or language), or sometimes by the cif/fob price ratio[3]. Because recent studies have shown that these variables are not the

[1] Polak (1996) emphasizes that if one doesn't use a measure of the average distance between a country and its main partners as well as absolute distance, one will underestimate trade between faraway countries.

[2] If relative distance is not taken into account, the dummy supposed to reflect trade between members of an agreement will capture the bias.

[3] Baier and Bergstrand (2001) use the CIF/FOB 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 CIF/FOB data see Hummels (1999) and Limao and Venables (2001).

only determinants, we model the barrier-to-trade function, between countries i and j, as follows[4] :

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

with expected signs on coefficients in parenthesis:

$DIST_{ij}$: distance between the countries i and j ($\delta_1>0$);

L_{ij} : takes the value 1 if i and j share a common border, otherwise 0 ($\delta_4<0$);

$E_{i(j)}$: takes the value 1 if the country i (j) is landlocked; otherwise 0 ($\delta_5>0$, $\delta_6>0$);

$IN_{i\,(j)}$: 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 ($\delta_2<0$, $\delta_3<0$).

Third, in a sample with countries that have large differences in income per capita, it is customary to abandon the homothetic utility function and allow Engel effects which implies including per capita income in the importing country and hence population N_i. On the supply side, it is reasonable to assume that supply will be driven by factor endowment differences. Following tradition, we use income per capita as proxy so that population in the exporting country, N_j is introduced in the model (e.g., Bergstrand, 1989, Frankel, 1997 or Soloaga and Winters, 2001).

Hence, after taking into account the modifications discussed above, the reduced form of the model is, after substitution of (3) in (1):

(5) $$\ln M_{ij} = \beta_0 + \beta_1 \ln Y_i + \beta_2 \ln Y_j + \beta_3 \ln N_i + \beta_4 \ln N_j + \beta_5 \ln DIST_{ij} + \beta_6 \ln \overline{DIST_i} + \beta_7 L_{ij}$$
$$+ \beta_8 E_i + \beta_9 \ln IN_i + \beta_{10} E_j + \beta_{11} \ln IN_j + \eta_{ij}$$

where Y^w is absorbed in the constant term, and with expected signs:

$\beta_1>0$, $\beta_2>0$, $\beta_3<0$, $\beta_4<0$, $\beta_5=-\sigma.\delta_1<0$, $\beta_6>0$, $\beta_7=-\sigma.\delta_4>0$, $\beta_8=-\sigma.\delta_5<0$, $\beta_9=-\sigma.\delta_2>0$,

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

β_{10}=-$\sigma.\delta_6$<0, β_{11}=-$\sigma.\delta_3$>0, and η_{ij} the error term.

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

First used by Aitken (1973) as an ex-post assessment for the EEC, the gravity model seems well-defined for this issue for two reasons. First, arguably, the model represents a relevant counterfactual (or anti-monde) 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.

Second, 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 an RTA, TC and TD will be reflected in trade flows as follows :(i) under pure TC intra-regional trade increases and imports from the ROW remains unchanged; (ii) under pure TD, the increase in intra-regional trade is entirely offset by a corresponding decrease of 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 an RTA for members. Finally, for non-members, because under plausible assumption about the anti-monde a necessary condition for their welfare to increase is that the volume of their imports increases once the RTA has been established (see Winters, 1997), one should include measure the change in volume of exports from members to non-members (an increase signifying an improvement in welfare for non-members).

Therefore, the correct ex-post assessment of an RTA on the volume of trade should include the following dummy variables (associated coefficients in parenthesis)[5] :

(i) D_I (α_I) =1 if both partners belong to the same RTA [zero otherwise] (captures 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] (captures 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] (captures bloc exports to the ROW).

Suppose that $\alpha_I > 0$ which corresponds to more intra-bloc trade than predicted by the reference ($\alpha_I < 0$ corresponding to an RTA between complementary economies) 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.

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[5] This is the specification in Endoh (1999) and Soloaga and Winters (2001). Others assessing the effects of RTAs (e.g., Bayoumi and Eichengreen, 1997, Frankel, 1997, Krueger, 1999) have not included enough dummy variables to distinguish between exports and imports, and hence fail to isolate TD and TC effects.

To summarize, following an 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 1962-96. Trade data are from UN COMTRADE (bilateral imports in current dollars). The dependent variable is total bilateral imports and is deflated by a world import price index taken from International Financial Statistics (IFS). Data sources for the explanatory variables along with data transformations are presented in appendix A.3. Once the missing values are taken out[6], the sample covers 130 countries (a list of the countries in the sample is presented in appendix A.4). There are thus 240 691 observations for 14 387 pairs of countries.

3.1 Panel specification

The usefulness of the gravity model to assess RTAs rests upon a plausible estimation of the anti-monde. 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.

[6] The countries which do not declare their imports from a partner or which do not import from this partner are identified in the same way, with a missing value. Hence, our data are not censored at zero. The actual number of observations (240 691) represent around 50% of potential number. The selection bias which can exist is tested and corrected by the inclusion of the selection rule in the model estimation in section 4.

historical, cultural, ethnic, political or geographical factors)[7] which is troublesome since the dummy variables should really isolate TD and TC effects. Not taking into account of countries' heterogeneity or of the pair of countries in bilateral trade relations introduces 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 inclusion of this bilateral term, α_{ij}, specific to each pair of countries and common to each year (and different according to the direction of trade: $\alpha_{ij} \neq \alpha_{ji}$), is more general than dummies capturing specific elements of trade such as common language or cultural similarity (see Cheng and Wall, 1999, Egger and Pfaffermayr, 2000).

So the previous model is specified in panel as:

$$(6) \qquad \ln M_{ijt} = \boldsymbol{\alpha_0 + \alpha_t + \alpha_{ij}} + \beta_1 \ln Y_{it} + \beta_2 \ln Y_{jt} + \beta_3 \ln N_{it} + \beta_4 \ln N_{jt} + \beta_5 \ln DIST_{ij}$$

$$+ \beta_6 \ln \overline{DIST_i} + \beta_7 L_{ij} + \beta_8 E_i + \beta_9 \ln IN_{it} + \beta_{10} E_j + \beta_{11} \ln IN_{jt} + \beta_{12} \ln RER_{ijt} + \eta'_{ijt}$$

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[8];

α_{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 (6). In a model with panel data that spans a long time period (here 35 years), it is essential to capture the evolution of competitiveness. Given our definition of e_{ij} in equation (1), an increase of the RER reflects a depreciation of the importing country's currency against that of the exporting country which should reduce imports (hence one would expect $\beta_{12}<0$).

[7] 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).

[8] These time dummies capture common shocks such as the evolution of oil prices over the period or Y^w in (1).

3.2 Econometric method

Since a fixed effects model is inadequate (the within transformation eliminates time-invariant variables), bilateral effects are modeled as random variables. Time effects are captured by yearly dummies that capture common shocks (e.g. oil price changes). 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 GDP or infrastructure may be correlated with bilateral specific effects[9]..

The usual way to deal with this issue is to consider an instrumental variables estimation such as that proposed by Hausman and Taylor (1981) (see appendix A.5 for the implementation of this method), though here it is adapted to the case of an unbalanced sample according to the method proposed by Guillotin and Sevestre (1994).

A Hausman-Taylor test of over-identification, based on the comparison of the Hausman-Taylor estimator and the Within one, must be carried out (see appendix A.5). 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 Hausman-Taylor estimator (HT) is the most efficient estimator.

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

3.3 Endogenity of explanatory variables and sample selection bias

Before proceeding to the evaluation of the effects of RTA (which are, in this section, captured by the bilateral specific effect), I check first for endogeneity of explanatory variables.[10] 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 a 99% 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 estimating the error component model (GLS) which differ markedly from the Within estimation. The Hausman test, based on differences between Within and GLS estimators, reveals a χ^2_7 = 462.07, which is significant at 99%. 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 Hausman-Taylor method justified.

For sensitivity analysis, four regressions are estimated with the Hausman-Taylor method. The over-identification test indicates for each regression if the instruments are legitimate or if an additional source of correlation between specific effects and explanatory variables exists (in the case of a significant test statistic).

[10] 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.

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)**
ln Y_{it}	1.03**	0.79**	1.00**	1.10**	1.19**	1.03**
	(49.8)	*(78.4)*	*(60.4)*	*(74.5)*	*(74.4)*	*(68.4)*
ln Y_{jt}	1.11**	1.12**	1.18**	1.17**	1.13**	1.11**
	(59.8)	*(129.8)*	*(109.4)*	*(107.0)*	*(98.4)*	*(80.2)*
ln N_{it}	0.19**	0.017	-0.088**	0.20**	0.11**	0.20**
	(5.15)	*(1.5)*	*(-6.6)*	*(8.9)*	*(4.7)*	*(7.7)*
ln N_{jt}	-0.65**	-0.19**	-0.25**	-0.63**	-0.65**	-0.65**
	(-24.8)	*(-17.4)*	*(-19.5)*	*(-24.1)*	*(-24.6)*	*(-24.6)*
ln $DIST_{ij}$	-	-1.14**	-1.17**	-1.19**	-2.09**	-1.19**
		(-59.0)	*(-51.1)*	*(-48.4)*	*(-14.3)*	*(-45.3)*
ln $\overline{DIST_i}$	-	-0.66**	0.26**	1.13**	2.02**	1.10**
		(-26.4)	*(5.1)*	*(17.1)*	*(22.2)*	*(15.8)*
L_{ij}	-	0.68**	1.14**	1.01**	0.77*	1.04**
		(9.6)	*(10.4)*	*(8.8)*	*(2.5)*	*(8.5)*
E_i	-	-0.27**	-0.18**	-0.02	-0.10	-0.02
		(-5.4)	*(-3.1)*	*(-0.4)*	*(-1.7)*	*(-0.3)*
E_i	-	-0.47**	-0.41**	-0.58**	-0.59**	-0.56**
		(-10.4)	*(-8.1)*	*(-11.2)*	*(-10.6)*	*(-9.6)*
ln IN_{it}	0.04**	0.04**	0.04**	0.05**	0.07**	0.04**
	(5.3)	*(6.3)*	*(6.8)*	*(6.3)*	*(9.9)*	*(6.0)*
ln IN_{jt}	0.03**	0.02**	0.01**	0.03**	0.03**	0.03**
	(5.9)	*(5.6)*	*(3.1)*	*(6.0)*	*(5.9)*	*(6.5)*
ln RER_{ijt}	-0.006**	-0.005**	-0.004**	-0.006*	-0.006**	-0.006**
	(-5.7)	*(-3.6)*	*(-2.7)*	*(-4.1)*	*(-4.4)*	*(-4.2)*
Number of obs (NT)	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
R^2 e)	0.87	0.63	0.62	0.61	0.61	0.63
Theta (mean)	-	0.82	0.83	0.83	0.83	0.84
Bilateral fixed effect	37.09**	-	-	-		-
	F(14386,226263)					
Time fixed effect	59.97**	161.31**	98.61**	131.58**	144.44**	156.31**
	F(34,226263)	*F(34,240644)*	*F(34,240644)*	*F(34,240644)*	*F(34,240644)*	*F(34,240644)*
Hausman test W vs. GLS f)	-	469.07**	-	-		-
chi-2(Kw)		*chi-2 (7)*				
Hausman test HT vs. GLS g)	-	-	961.27**	1184.41**	1480.83**	973.53**
chi-2(K)			*chi-2(12)*	*chi-2(12)*	*chi-2(12)*	*chi-2(12)*
Test of over-identification h)	-	-	243.91**	13.21**	110.33**	0.39
chi-2 (k_1-g_2)			*chi-2(5)*	*chi-2(3)*	*chi-2(1)*	*chi-2(1)*

** and * significant at 99% and 95% 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 = $\ln Y_{it}$ et $\ln Y_{jt}$, k_1-g_2=5.

b) HT II : endogenous variables = $\ln Y_{it}$, $\ln Y_{jt}$, $\ln N_{it}$ et $\ln N_{jt}$, k_1-g_2=3.

c) HT III : endogenous variables =$\ln Y_{it}$, $\ln Y_{jt}$, $\ln N_{it}$, $\ln N_{jt}$, $\ln DIST_{ij}$ et $\ln \overline{DIST_i}$, k_1-g_2=1.

d) HT IV : endogenous variables = $\ln Y_{it}$, $\ln Y_{jt}$, $\ln N_{it}$, $\ln N_{jt}$, $\ln IN_{it}$ et $\ln IN_{jt}$, k_1-g_2=1.

e) 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.

f) 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(41)= 853.61**

g) Hausman test applied to the differences between GLS and HT estimators, without time effects.

h) Hausman test applied to the differences between Within and HT estimators, without time effects.
Cf. a), b) c) and d) for information on k_1-g_2.

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 I to GLS, confirms that the instrumentation has improved the model[11] (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 (χ^2_5 = 243.91). 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 can be due to the variables of infrastructure[12]. Their instrumentation, in addition to those of 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 (see appendix A.5), the Hausman-Taylor estimator is convergent and more efficient than the Within estimator[13].

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[11] Guillotin and Sevestre (1994) recommend comparing the HT estimator, denoted β_{HT}, to the GLS estimator, denoted β_{MCQG}. It is exactly the same principle as for earlier tests presented in appendix A.5. We compute the Hausman statistic test and the number obtained is compared to the critical value of χ_K^2 (K is the dimension of the coefficients' vector β_{MCQG}). 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, the instrumented variables are actually endogenous.

[12] I also test for the correlation of distance variables with bilateral effects in the equation HT III. However this equation does not improve the model HT II.

[13] According to the Barghava and al. Durbin Watson test (1982), modified to the unbalanced panel, the HT IV residuals are no autocorrelated AR(1): there is no systematic difference between observed and predicted trade flows. Hence, the HT IV estimator is efficient and the over-identification test is appropriate (e.g. Egger 2002).

All these coefficients are significant at a 99% level (except for E_i) and have the expected sign. Import volume of i from j increases with GDP and coefficients are close to unity as suggested by the theory and as reported by, for instance, Aitken (1973), Soloaga and Winters (2001) and Egger and Pfaffermayr (2000). The population variable has the expected negative sign for the exporting country (capturing the fact that larger countries trade less) but has a positive sign for the importing country (as in e.g., Soloaga and Winters, 2001, and Egger and Pfaffermayr, 2000). The elasticity of bilateral trade to distance is superior to unity (-1.19) and 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 2.8 times more than expected from the gravity equation (=exp(1.04)). Likewise, imports from a country without direct access to the sea are 43% lower. Finally, a real depreciation of i with respect to j lowers i's imports from j.

A last potential estimation bias must be considered: the unbalanced sample can be subject to a non-ignorable selection rule, 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. 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[14], by adding variables which reflect the individual's patterns in terms of presence in the sample to the model. 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)

[14] I use an alternative method because the generalization of the Heckman two-steps method (1979) to panel data and random effects model is too difficult (Guillotin and Sevestre 1994, p.127).

DD: dummy that takes 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 appendix A.6 and are compared to equation HT IV (table 1). In the first column, with only the variable PRES considered, the conclusions of the previous estimations are not modified, even if the coefficient of PRES is statistically different from zero. The following regressions (columns 2 and 3) show that the variables DD and PA_t have a positive and significant coefficient: all other things equal, pairs of countries which have at least two years of consecutive available data (and *a fortiori* if they are present over the entire period) 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

Following the specification check, the three dummy variables discussed above were introduced in the model to detect TD and TC for a selection of RTAs (EU, ANDEAN, NAFTA, CACM, MERCOSUR, ASEAN, EFTA, LAIA). To save space, detailed comments are only reported for three well-known RTAs: EU, NAFTA, and MERCOSUR, the EU being included in spite of lack of data for years prior to the agreement because it is the best-known and most studied RTA. Average effects over the sample period are reported first, then effects over time to look for break points around the important dates of the agreements.

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. All results are presented in appendix A.7.

Table 2 : Results for regional dummies over 1962-96.

Variables	M_{ijt}	
	Panel (HT IV)	**Cross-section (average coefficients) a)**
EU_{intra}	0.291*	-0.215
$EU_{imports}$	0.225**	0.797
$EU_{exports}$	0.375**	0.746
$MERCOSUR_{intra}$	-0.275	-0.432
$MERCOSUR_{imports}$	-1.041**	0.017
$MERCOSUR_{exports}$	-0.130**	0.088
$NAFTA_{intra}$	-0.063	0.754
$NAFTA_{imports}$	-0.478**	0.253
$NAFTA_{exports}$	0.009	0.011

** and * significant at 99% and 95% respectively for Panel estimation.
a) For each variable, this is the average of the 35 coefficients estimated per year from 1962 to 1996.

Generally, the significance of the coefficients is greater for the panel specification, with coefficients of the same sign when they are significant in both specifications.

These result give the average impact of each RTA over 1962-96. However, relevant inferences about TD and TC require inspection of the evolution of these coefficients over time and around the period when RTAs go into effect which can be done by breaking down estimation into subperiods. To this effect, I break down regional dummy variables into two-year periods with these variables introduced in the estimating equation instead of the global regional dummies.[15]

[15] Coefficient estimates for explanatory variables are identical under this procedure because for each agreement, the addition of the new variables introduced is equal to the former aggregate dummy variable.

4.2 Evolution of the effects during the RA's existence

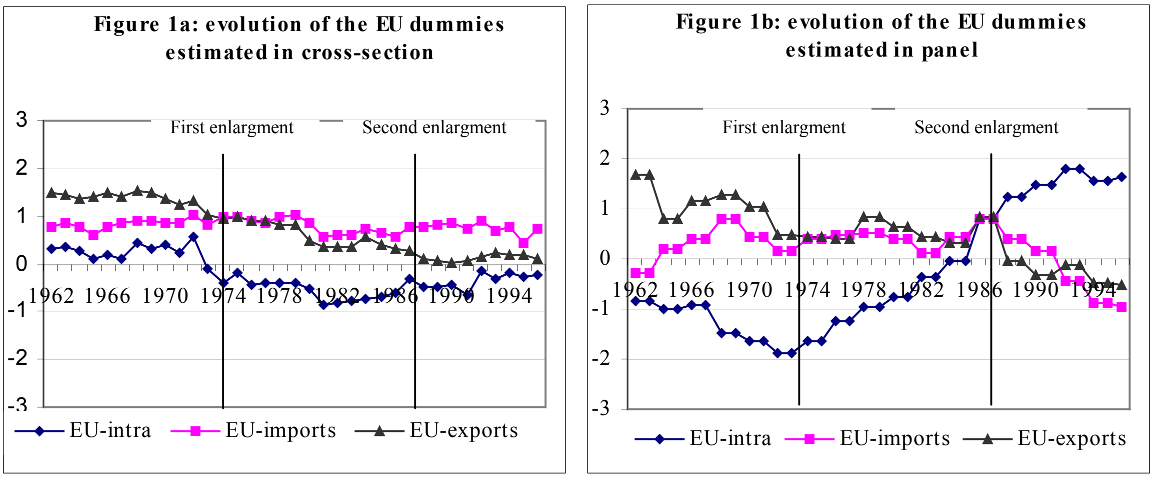
Because the results are self-explanatory from inspection of figure 1 to 3, I comment the EU results and give only an overall interpretation for NAFTA and MERCOSUR along with a summary for other RTAs[16]. All the evolutions commented in this section are significant one[17].

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

[16] The evolution of the estimated coefficients are represented in appendix A.8 for cross-section results and for panel ones for the other RTAs.

[17] The consecutive coefficients are tested significantly different.

Figure 1: evolution of EU dummies over 1962-1996 (α_I, α_M and α_X)



By contrast, panel estimates (figure 1b) suggest three rather distinct periods in terms of TC and TD. From 1967 to 1973, intra-trade decreases somewhat surprisingly without clear tendencies for trade with ROW. However, following the first (and second) enlargements, the models predicts a significant positive trend in intra-trade (α_I increases and turns positive in 1984, the pattern continuing with the deep integration following the EC-92 programme). In parallel, there is first a stagnation of imports of 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) who found a TD after the first enlargement and TC after the second, but are closer to those of Frankel (1997), Krueger (1999) or Soloaga and Winters (2001).

Figure 2: evolution of MERCOSUR dummies over 1962-1996 (α_I, α_M and α_X)

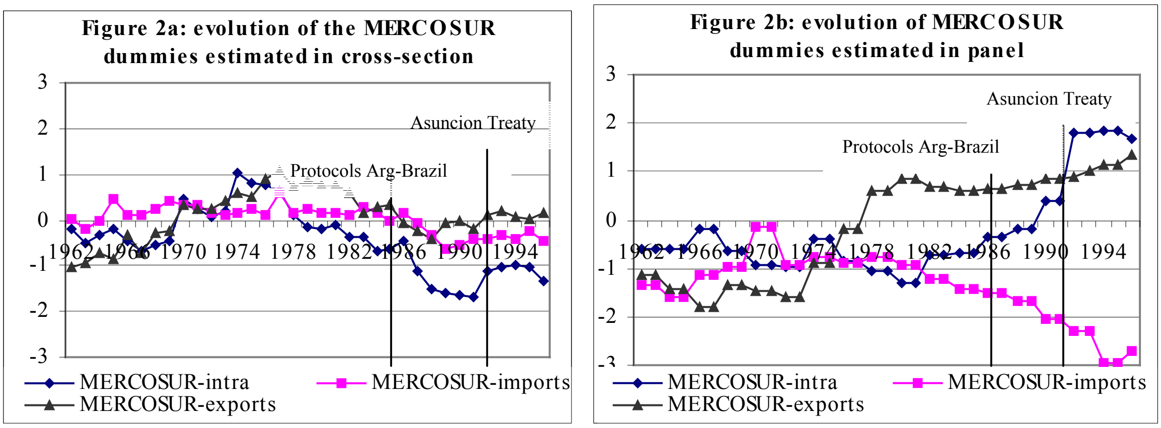
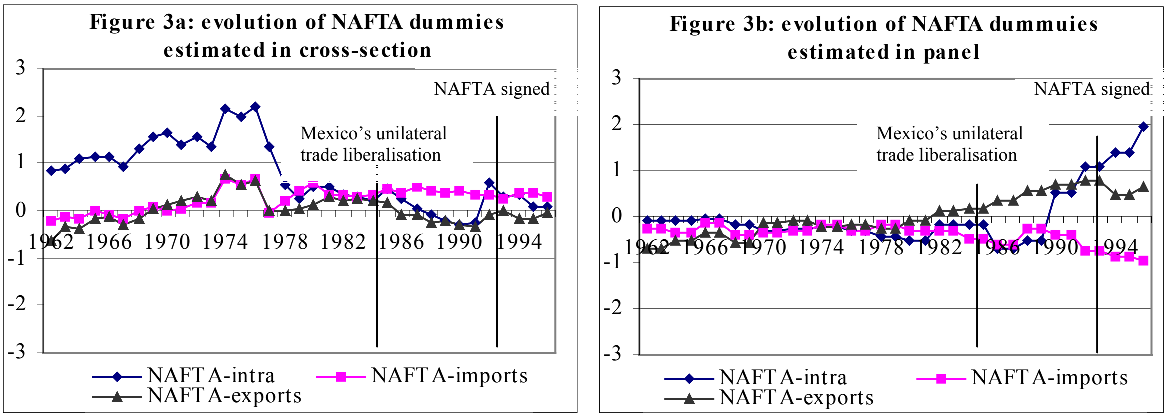


Figure 3: evolution of NAFTA dummies over 1962-1996 (α_I, α_M and α_X)



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 an RTA around the time of announcement or 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 the exports for NAFTA and MERCOSUR to the ROW (which probably reflects the opening up of the countries to the world as the same time as they were forming the RTA). Clearly, the panel estimates reveal a more plausible pattern than the cross-section estimates.

This pattern of import (and sometimes export) TD was also found for other RTAs reported in appendix A.8. For example, in the case of the ANDEAN accord, the model finds import-TD over the period 1969-79, over the period 1962-77 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 ROW for CACM. No clear patterns emerged for EFTA, while ASEAN and LAIA 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 correction for endogeneity of the income, size and infrastructure 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 anti-monde 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 examination of three well-studied RTAs: EU, MERCOSUR and NAFTA. In general, the results in this study, covering eight RTAs, show that most of them resulted in an increase in intra-regional trade beyond levels predicted by the anti-monde reference, often coupled with a reduction in imports from the ROW, and at times coupled with a reduction in exports to the ROW, suggesting evidence of trade diversion.

Appendices

A.1 : Derivation of the gravity model

As in Deardorff (1998), assume each country i is specialized in a single commodity, with a representative consumer maximizing a homothetic utility function:

$$U_i = \left(\sum_j b_j C_{ji}^{\left(\frac{\sigma-1}{\sigma}\right)} \right)^{\frac{\sigma}{\sigma-1}} \qquad \text{(A1)}$$

where σ is the common elasticity of substitution between any pair of countries' products subject (σ>0), and $b_j=b_i$, ∀i,j guarantees symmetry and a single price for each product variety. Product differentiation is at the national level (rather than at the firm level as in the monopolistic competition version), and CES preferences (rather than Cobb-Douglas) implies that bilateral trade decreases with distance. Each consumer Maximization of (A1) subject to the budget constraint $Y_i=p_i x_i$ (with x_i the production of country i) gives:

$$C_{ji} = \frac{1}{p_i} b_j \left(\frac{p_i}{\overline{P}_i} \right)^{1-\sigma} Y_i \qquad \text{(A2)}$$

where $$\overline{P}_i = \left(\sum_j b_j p_i^{1-\sigma} \right)^{1/(1-\sigma)} \qquad \text{(A3)}$$

is the CES price aggregator in country i associated with the minimization of expenditures in the utility maximization problem and p_i is the price in the country of destination i facing consumers. Assume that the relationship between the price in the country of origin j, p_j, and the country of destination i, p_i is given by :

$$p_i = p_j e_{ij} \theta_{ij} \qquad \text{(A4)}$$

In (A4), e_{ij} represents the nominal bilateral exchange rate and θ_{ij} the barrier-to-trade function between i and j. This term is usually proxy by the distance between the two countries.

To get the standard gravity-based model, assume balanced trade and let $\gamma_j=Y_j/Y^W$ be the share of country j in world income, Y^W. Expenditures of all countries i on the good produced in j are $\sum_i p_i C_{ji}$. Then, $Y_j=\sum_i p_i C_{ji}$ and substituting the value of C_{ji} from (2) into this expression gives:

$$b_j=\gamma_j\left(\sum_i \gamma_i\left(\frac{p_i}{\overline{P}_i}\right)^{1-\sigma}\right)^{-1} \tag{A5}$$

Substituting (A5) into (A2), the volume of imports of country i from j is given by:

$$M_{ji}=\frac{Y_iY_j}{Y^W}\left[\frac{\dfrac{(p_i)^{-\sigma}}{(\overline{P}_i)^{1-\sigma}}}{\sum_h \gamma_h\left(\dfrac{p_i}{\overline{P}_h}\right)^{1-\sigma}}\right]=\frac{Y_iY_j}{Y^W}\theta_{ij}^{-\sigma}e_{ij}^{-\sigma}\left[\frac{\dfrac{(p_j)^{-\sigma}}{(\overline{P}_i)^{1-\sigma}}}{\sum_h \gamma_h\left(\dfrac{p_i}{\overline{P}_h}\right)^{1-\sigma}}\right] \qquad \forall i,j,h=1..n \tag{A6}$$

The intensity of trade between two countries is a function of their respective size and that it is a decreasing function of the extent of barriers to trade θ_{ij}.

To simplify this, first select units of goods so that each country's product price, p_j, is normalized to unity (and e_{ij}=1). Then, as shown by Deardorff, $\overline{P}_i$ (given by A3) becomes a CES index of country i's barriers-to-trade factors as an importer. Using Deardorff's notation, the average barrier-to-trade from suppliers, δ_i^S, is given by:

$$\delta_i^S=\left(\sum_j b_j(\theta_{ij})^{1-\sigma}\right)^{1/(1-\sigma)} \tag{A7}$$

Substituting (A7) into (A6) gives expression:

$$M_{ji}=\frac{Y_iY_j}{Y_W}\theta_{ij}^{-\sigma}\left[\frac{\left(\dfrac{1}{\delta_i^S}\right)^{1-\sigma}}{\sum_h \gamma_h\left(\dfrac{\theta_{hj}}{\delta_h^S}\right)^{1-\sigma}}\right] \tag{A8}$$

A.2 : Definition of the regional agreements studied

	UE	EFTA	NAFTA	LAIA	CACM	ANDEAN	MERCO SUR	ASEAN
1962	**1957**(EEC)	**1960**		**1960** (LAFTA)	**1960**			
	France	Austria		Argentina	Costa Rica			
	Germany	Denmark		Bolivia	El Salvador			
	Belgium	Norway		Brazil	Guatemala			
	Italy	Portugal		Chile	Honduras			
	Luxembourg	Sweden		Colombia	Nicaragua			
1964	Netherlands	Switzerland		Ecuador				
		UK		Mexico				
		Finland		Paraguay				
1966				Peru				
				Uruguay				**1967**
				Venezuela				Indonesia
1969						**1969**		Singapore
						Bolivia		Philippines
						Chile		Malaysia
		1970				Colombia		Thailand
1973	**1973**(EEC)	Austria				Ecuador		
	France	Iceland(70)				Peru		
1975	Germany	Norway				Venezuela(73)		
	Belgium	Portugal				**1976**		
	Italy	Sweden				Bolivia		
	Luxembourg	Switzerland				Colombia		
1980	Netherlands	Finland		**1980** (LAIA)		Ecuador		
	UK			Argentina		Peru		
	Denmark			Bolivia		Venezuela		
1984	Ireland			Brazil				
	Greece (81)	**1985**		Chile				
	Spain (86)	Austria		Colombia				
	Portugal (86)	Iceland(70)		Ecuador				
	Austria (95)	Norway		Mexico				
	Finland (95)	Sweden		Paraguay				
	Sweden (95)	Switzerland		Peru				
		Finland		Uruguay				
1991				Venezuela			**1991**	
							Argentina	
1992			**1992**				Brazil	**1992**
			Canada			**1992**	Uruguay	Indonesia
			Mexico			Bolivia	Paraguay	Singapore
1994			USA			Colombia		Philippines
						Ecuador		Malaysia
						Venezuela		Thailand
		1995						
		Norway						
		Switzerland						
		Liechtenstein						
1996		*(91)*						

Bilateral trade of Liechtenstein and Switzerland is not desegregated in this data set (as for Belgium and Luxembourg).

A.3 : Sources and data definition

M_{ijt} : COMTRADE, total bilateral imports of country i from country j at time t. This variable is in current dollar so it has been divided by an index of the unit value of imports, which is taken from IMF, to obtain a real flow of trade.

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

$N_{i(j)t}$: CD-ROM WDI, World Bank 1999, total population of country i at time t.

$DIST_{ij}$: Data for distance are extracted from the software developed by the company CVN. The distance is measured in kilometers between the main city of the country i and that of country j. Most of the time, the main city is the capital city, but for some countries the main economic city is considered. The distance calculated by this software is orthodromic, that is, it takes into account the sphericity of Earth. More precisely, 'the distance between two points A and B is measured by the arc of the circle subtended by the chord [AB]' (see HAINRY, «Jeux Mathématiques et Logiques – Orthodromie et Loxodromie »).

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

$E_{i(j)}$: Dummy equal to one if the country i is landlocked (i.e. do not have a direct access to the sea), 0 otherwise.

$IN_{i(j)t}$: This index is built using 4 variables from the database constructed by Canning (1996): the number of kilometer of roads, of paved roads, of railways, and the number of telephone sets/lines per capita of country i (j) at time t. The first three variables are divided by the land area (WB, 1999) to obtain a density. Thus, each variable obtained is normalized to have a same mean equal to one. An arithmetic average is then calculated over the four variables, for each country and each year, without taking into account the missing values (a similar computation is presented by Limao and Venables 2001). As the final year of the data set is 1995, an extrapolation had to be made to cover the year 1996.

$\overline{DIST_i}$: average distance of country i to exporter partners, weighted by exporters' GDP share in world GDP ("remoteness" of country i). The ten main trade partners are identified for each country according to bilateral flows averaged over 1980-96 (in COMTRADE). For the weights, we used 1990's GDP (WB, 1999). Hence, This variable is specific to each country and is not time variant.

RER_{ijt} : We extract from the IFS data set the nominal exchange rate for each country against US dollar ($NER_{i/\$}$, 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: $RER_{i/j} = (CPI_j) / (CPI_i) . (NER_{i/\$} / NER_{j/\$})$, where i is the importing country and j the exporting one. For each pair of countries, we specify the RER such as its mean over the period is zero.

A.4 : Countries in the sample.

OECD	Sub-Saharan Africa	Latin America and the Caribbean	Asia and the Pacific	Others
Australia	Angola	Argentina	Bangladesh	*Albania*
Austria	South Africa*	Bahamas	Brunei	*Armenia*
Belgium +	Burundi	Barbados	Bhutan	*Azerbaijan*
Luxembourg	Benin	Belize	China	Bulgaria
Canada	Burkina Faso	Bolivia	Fiji	*Belarus*
Germany	Central African Rep.	Brazil	Hong Kong	Czech Rep.
Denmark	Ivory Coast	Chile	Indonesia	Algeria
Spain	Cameroon	Colombia	India	Saudi Arabia
Finland	Congo	Costa Rica	*Cambodia*	Egypt
France	*Comoros*	Dominican Rep.	*Lao PDR*	Estonia
United Kingdom	Cape Verde	Dominica	Macao	*Georgia*
Ireland	Djibouti	Ecuador	*Mongolia*	Greece
Iceland	Ethiopia + Eritrea	Grenada	Malaysia	Bosnia and
Italy	Gabon	Guatemala	*Nepal*	Herzegovina
Japan	Ghana	Guyana	Pakistan	Hungary
Korea, Rep.	Guinea	Honduras	Philippines	Iran
United States	Guinea-Bissau	Haiti	Papua New Guinea	Israel
Netherlands	Gambia	Jamaica	Singapore	Jordan
Norway	Equatorial Guinea	Mexico	Salomon Islands	*Kazakstan*
New Zealand	Kenya	Nicaragua	Thailand	*Kyrgyz Rep.*
Portugal	Madagascar	Panama	*Vietnam*	Kuwait
Sweden	Mali	Peru	*Western Samoa*	Lithuania
Switzerland +	Mozambique	Paraguay	Sri Lanka	Latvia
Liechtenstein	Mauritania	El Salvador	*Tonga*	*Macedonia*
	Mauritius	Suriname	*Kiribati*	Morocco
	Malawi	Trinidad and Tobago	*Vanuatu*	Malta
	Niger	Uruguay		Oman
	Nigeria	St. Vincent and		Poland
	Rwanda	The Grenadines		Romania
	Sudan	Venezuela		*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.5: the Hausman and Taylor (1981) method.

Let us consider:

$$M_{ijt}= X_{ijt}\beta + Z_{ij}\delta+ u_{ijt} \quad \text{with } u_{ijt} = \alpha_{ij} + \nu_{ijt} \qquad (A.9)$$

With $X_{ijt} = [\ln Y_{it}\ \ln Y_{jt}\ \ln N_{it}\ \ln N_{jt}\ \ln IN_{it}\ \ln IN_{jt}\ \ln RER_{ijt}]$

and $Z_{ij}= [\ln DIST_{ij}\ \ln \overline{DIST_i}\ \ L_{ij}\ E_i\ E_j]$

where some explanatory variables of X (variables variant over time) and of Z (time-invariant variables) are correlated with the specific effects. We 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[18] and (A.9) can be estimated using $[QX_1, QX_2, PX_1, Z_1]$[19] as instruments (see Breusch, Mizon and Schmidt, 1989). The instruments are then taken within the model. The resulting estimator is consistent but not efficient, as it does not correct for heteroskedasticity and serial correlation due to the presence of random bilateral specific effects. Hence, Hausman and Taylor (1981) 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). The instrumental variable estimator is then applied to the following transformed equation:

$$Y_{ijt} - (1-\theta)\, Y_{ij.} = [\, X_{ijt} - (1-\theta)X_{ij.}\,]\, \beta + \theta Z_{ij}\delta + \theta\mu_{ij} + [\nu_{ijt} - (1-\theta)\, \nu_{ij.}]$$

With[20] $\theta= (\sigma_v^2 / T\sigma_\mu^2 + \sigma_v^2)^{1/2}$ *(A.10)*

A test of over-identification must be carried out. It is based on the comparison of the Hausman-Taylor estimator, denoted β_{HT}, and the Within estimator (fixed effects model), denoted β_w. The Hausman test statistic is:

$$[\beta_{HT} - \beta_W] \,.\, [var(\beta_{HT}) - var(\beta_W)]^{-1} \,.\, [\beta_{HT} - \beta_W]' \qquad (A.11)$$

Under the null hypothesis, this test statistic is distributed as a Chi-square (χ^2) with k_1-g_2 degrees of freedom. If the statistic is inferior to the critical value, then the null can't be rejected: the instruments are legitimate[21]. If $k_1 > g_2$, we can also conclude that the Hausman-Taylor estimator (HT) is the most efficient estimator.

---

[18] If $k_1 > g_2$ then the equation is over-identified.

[19] 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).

[20] Owing to the fact that our sample is unbalanced, we have in fact $\theta_{ij}= (\sigma_v^2 / T_{ij}\sigma_\mu^2 + \sigma_v^2)^{1/2}$ with σ_μ^2 et σ_v^2 which are corrected for the bias of heteroskedasticity, specific to the unbalanced sample, according to the method proposed by Guillotin and Sevestre (1994). The mean value of θ_{ij} will be systematically presented in the tables of results.

[21] Actually, the null hypothesis H_0 is that there are no significant difference between the Within estimator and the HT one. So, under H_0, there is no longer bias due to a correlation between specific bilateral effects and explanatory variables.

A.6 : Test and correction of selection bias in equation HT IV

Variables	M_{ijt}		
	1	2	3
ln Y_{it}	1.00**	1.06**	0.98**
	(61.1)	*(65.0)*	*(60.8)*
ln Y_{jt}	1.13**	1.15**	1.14**
	(77.1)	*(76.9)*	*(89.5)*
ln N_{it}	0.16**	0.19**	0.16**
	(6.4)	*(7.3)*	*(6.2)*
ln N_{jt}	-0.64**	-0.69**	-0.62**
	(-21.4)	*(-26.1)*	*(-22.8)*
ln $DIST_{ij}$	-1.09**	-1.14**	-1.17**
	(-44.8)	*(-43.8)*	*(-43.9)*
ln $\overline{DIST_i}$	0.96**	1.17**	0.60**
	(14.3)	*(17.0)*	*(8.7)*
L_{ij}	0.97**	1.04**	0.90**
	(8.8)	*(8.7)*	*(7.2)*
E_i	-0.17	-0.05	-0.14*
	(-5.2)	*(-0.8)*	*(-2.3)*
E_j	-0.54**	-0.52**	-0.51**
	(-6.4)	*(-9.1)*	*(-4.9)*
ln IN_{it}	0.04**	0.04**	0.04**
	(5.1)	*(5.6)*	*(4.6)*
ln IN_{jt}	0.03**	0.03**	0.03**
	(7.3)	*(6.6)*	*(7.1)*
ln RER_{ijt}	-0.006**	-0.006**	-0.005**
	(-4.1)	*(-4.3)*	*(-3.1)*
PRES	**0.05****	**-**	**0.039****
	(29.2)		***(18.5)***
DD	**-**	**0.84****	**0.10**
		(13.5)	***(1.5)***
PA_t	**-**		**0.49****
			(49.4)
Number of obs (NT)	240 691	240 691	240 691
Number of bilateral (N)	14 387	14 387	14 387
R^2	0.63	0.64	0.65
Theta (mean)	0.83	0.84	0.85

** and * significant at 99% and 95% respectively (t-student is presented under the correspondent coefficient).
The time dummy variables and the constant are not reported in order to save space.
The estimation method is one of Hausman-Taylor, with variables Y_{it}, Y_{jt}, N_{it}, N_{jt}, IN_{it} and IN_{jt} as endogenous (HT IV).

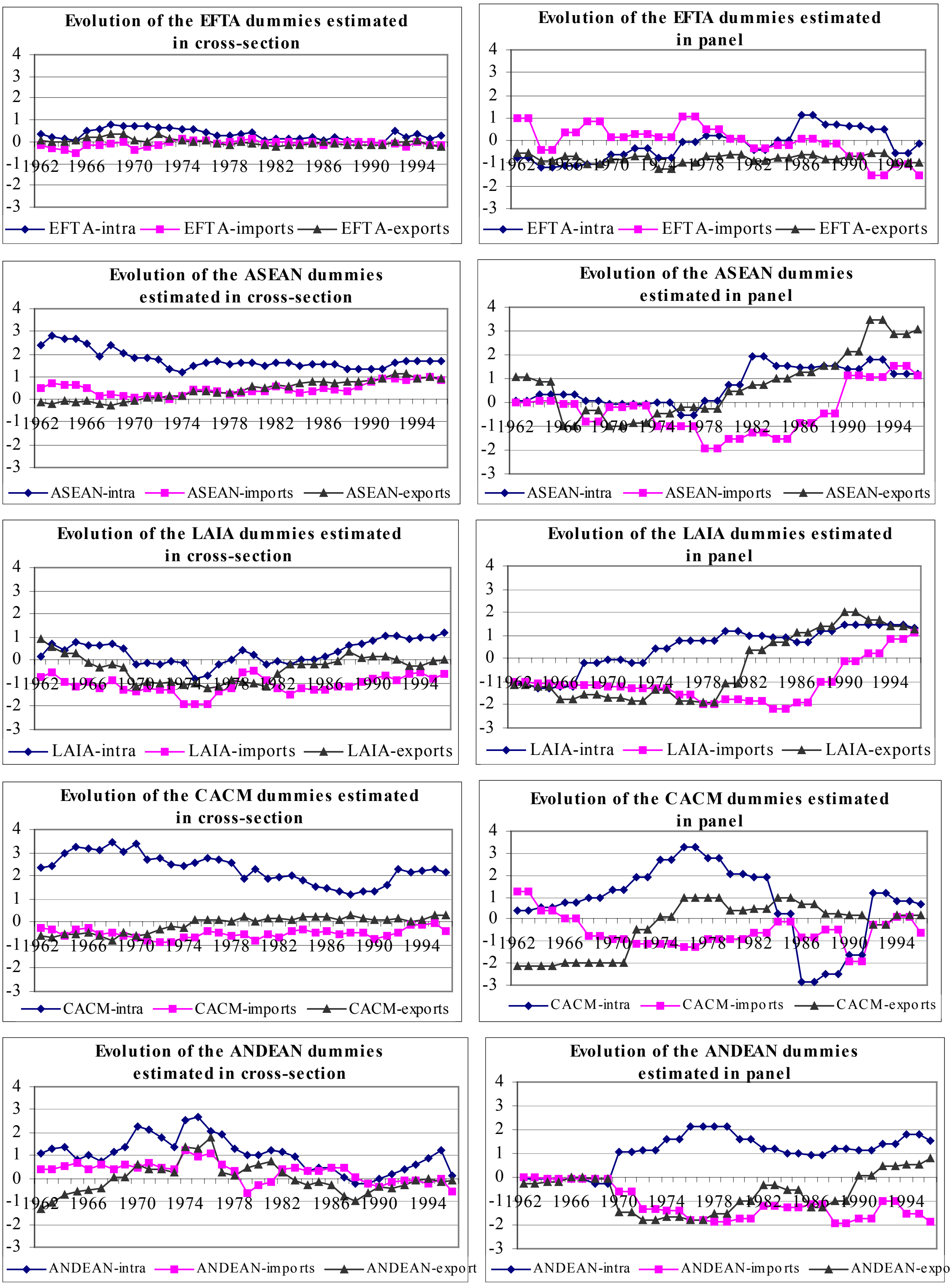
A.7 : Results of the estimation with regional dummies (1962-1996).

Variables	M_{ijt}				
	Panel		Cross-section		
	Coeff.	t	Average Coeff.	max	min
$\ln Y_{it}$	1.033**	*62.66*	0.769	*0.97*	*0.63*
$\ln Y_{jt}$	1.139**	*85.15*	0.935	*1.21*	*0.69*
$\ln N_{it}$	0.131**	*11.38*	-0.045	*-0.01*	*-0.12*
$\ln N_{jt}$	-0.650**	*-22.69*	-0.082	*-0.01*	*-0.20*
$\ln DIST_{ij}$	-1.168**	*-41.93*	-0.971	*-0.46*	*-1.25*
$\ln \overline{DIST_i}$	0.751**	*16.11*	0.136	*0.65*	*-0.71*
L_{ij}	0.921**	*7.68*	1.018	*1.58*	*0.54*
E_i	-0.161**	*-3.66*	-0.178	*-0.01*	*-0.63*
E_j	-0.510*	*-2.09*	-0.389	*-0.05*	*-1.07*
$\ln IN_{it}$	0.037**	*4.68*	0.157	*0.26*	*0.07*
$\ln IN_{jt}$	0.031**	*7.19*	0.067	*0.14*	*0.01*
$\ln RER_{ijt}$	-0.005**	*-4.12*	-	-	-
PRES	0.039**	*19.21*	-	-	-
DD	0.099	*0.31*	-	-	-
PA_t	0.494**	*49.4*	-	-	-
EU $_{intra}$	0.291*	*1.98*	-0.215	*0.58*	*-0.88*
EU $_{imports}$	0.225**	*3.16*	0.797	*1.04*	*0.04*
EU $_{exports}$	0.375**	*5.11*	0.746	*1.53*	*0.02*
EFTA $_{intra}$	-0.287	*-1.56*	0.319	*0.77*	*-0.10*
EFTA $_{imports}$	-0.075	*-0.89*	-0.098	*0.12*	*-0.55*
EFTA $_{exports}$	-0.932**	*-11.93*	-0.007	*0.37*	*-0.21*
ASEAN $_{intra}$	0.680**	*4.18*	1.757	*2.78*	*1.22*
ASEAN $_{imports}$	-0.513**	*-5.77*	0.458	*0.96*	*0.01*
ASEAN $_{exports}$	0.757**	*8.86*	0.421	*1.15*	*-0.25*
ANDEAN $_{intra}$	0.772**	*4.78*	1.049	*2.65*	*-0.25*
ANDEAN $_{imports}$	-0.940**	*-5.07*	0.285	*1.2*	*-0.64*
ANDEAN $_{exports}$	-0.959**	*-6.08*	-0.022	*1.79*	*-1.37*
MERCOSUR $_{intra}$	-0.275	*1.54*	-0.432	*1.01*	*-1.68*
MERCOSUR $_{imports}$	-1.041**	*-6.76*	0.017	*0.57*	*-0.63*
MERCOSUR $_{exports}$	-0.130	*0.93*	0.088	*1.06*	*-1.02*
LAIA $_{intra}$ **a)**	0.360**	*4.62*	0.327	*1.2*	*-0.77*
LAIA $_{imports}$	-1.492**	*-12.23*	-1.073	*-0.48*	*-1.92*
LAIA $_{exports}$	-0.357	*1.41*	-0.359	*0.88*	*-1.23*
CACM $_{intra}$	1.087**	*3.91*	2.305	*3.44*	*1.19*
CACM $_{imports}$	-0.776**	*-8.30*	-0.498	*0.06*	*-0.88*
CACM $_{exports}$	-0.127	*-1.34*	-0.097	*0.28*	*-0.79*
NAFTA $_{intra}$	-0.063	*-0.48*	0.754	*2.18*	*-0.30*
NAFTA $_{imports}$	-0.478**	*-5.96*	0.253	*0.68*	*-0.21*
NAFTA $_{exports}$	0.009	*0.07*	0.011	*0.76*	*-0.65*
Number of obs (NT)	240 691		7 265	*9 362*	*5 819*
Number of bilateral (N)	14 387		-	-	-
R^2	0.66		0.64	*0.73*	*0.60*
Theta (mean)	0.84		-	-	-

** and * significant at 99% and 95% respectively (t-student is presented next to 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)) :
intra-LAIA=LAIA-ANDEAN-MERCOSUR
LAIA imports= LAIA imports-ANDEAN imports-MERCOSUR imports
LAIA exports= LAIA exports-ANDEAN exports-MERCOSUR exports.

A.8 : Evolution of the RTA dummies estimated in panel and in cross-section over 1962-1996 (α_I, α_M and α_X).



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