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A survey of the assessments of the effectiveness of Preferential Trade Agreements using gravity models

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A survey of the assessments of the effectiveness of Preferential Trade Agreements using gravity models*

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Abstract

This paper reviews the empirical literature using a gravity model to assess the impact on trade of Preferential Trade Agreements (PTAs) and provides a critical analysis of the choices made to carry out the estimations. With respect to the effectiveness of PTAs, the main result that emerges from the review is that PTAs tend to foster trade between the countries involved. However, two empirical issues appear relevant. First, the use of dummy variables to proxy PTAs does not adequately describe the preferences granted and can be misleading. Indeed, the use of dummies does not allow to discern among the different preferential trade policy *instruments* used as well as to measure the *level* of the trade preferences granted. As a result, dummies do not allow to capture the specific effect of the preferences on bilateral trade. Second, the econometric methods used do not always address in a satisfactory way the potential sources of bias in the estimations, such as unobserved heterogeneity, endogeneity of some regressors and zero-trade flows. Thus, with respect to the empirical issues, the main conclusion which can be drawn from this survey is that the proposed estimates so far of the effects of PTAs using gravity models tend to be biased and, as a result, are not fully reliable.

1. Introduction

In the last decade there has been a growing interest in the effectiveness of Preferential Trade Agreements (PTAs), likely due to their proliferation. Indeed, in 2006 there were over 190 PTAs notified to the World Trade Organization (WTO). The main question asked is whether PTAs foster trade among those countries who participate in them. Many studies apply a gravity specification to answer this question. In its basic form, the gravity equation predicts trade flows as a function of the size of the trade partners and the distance between them. The original specification proposed by Tinbergern (1962) and Pöyhonen (1963) did not have a theoretical foundation. Since then, many offered a theoretical background for the gravity specification (Anderson, 1979; Bergstrand, 1985 and 1989; Deardoff, 1995; Anderson and

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van Wincoop, 2003); however, these attempts "have used ad hoc arguments to explain its validity, which makes every result potentially disputable on economic grounds" (Verdeja, 2006, p. 2).

In this survey papers which assess the impact of PTAs on trade using gravity models are reviewed and a critical analysis of the estimation approaches chosen is provided. The review covers over than 115 contributions, that is all the studies which, to the best of my knowledge, deal with this issue. A similar review was carried out by Greenaway and Milner (2002), who focused on the studies covering major Regional Trade Agreements (RTAs) in Europe, North America and Asia and considered about twenty contributions. Another paper aiming at summarizing the analyses of the role of PTAs on trade using gravity models is that of Cipollina and Salvatici (2006). However, their goal is quite different from that of this paper as they implemented a meta-analysis, that is an empirical method allowing to evaluate and synthesize results from different studies. Finally, Nielsen (2003) reviewed papers assessing the impact of PTAs in favour of developing countries and provided an overview of the empirical methods (partial equilibrium, general equilibrium and gravity models) used to analyse PTAs.

The main results of this survey can be summarized as follows. First, all the reviewed papers but two use a dummy variable to proxy the PTA. This dummy is equal to one if both trading partners belong to a PTA and zero otherwise. However, the use of a dummy to capture the impact of a PTA on trade is not adequate because it also captures all other factors that are specific to the country-pair and contemporaneous to the PTA. The only exceptions to the use of a dummy to proxy PTAs are provided by Cipollina and Salvatici (2007), who compute preferential margins, and by Emlinger et al. (2006), who consider explicitly preferential tariffs applied by the EU on its imports from partners.

Second, in the studies considered the results obtained with respect to the trade impact of PTAs are very heterogeneous, both in size and in sign. This could be due to the fact that different countries and periods are used and also the set of variables considered in the gravity specification is not homogeneous. For example, some papers consider only standard gravity variables, such as GDPs, populations and distance, while others also consider the level of infrastructure in both trading partners, the exchange rate and the exchange rate volatility.

Third, the estimation methods implemented are also highly diversified; all studies reviewed seem to disregard one or more estimation issues, leading to biased results. [1] In fact,

[1] Also Nielsen (2003) pointed out that some empirical issues (zero-trade flows and heterogeneity) are not always properly considered in studies using gravity model to assess the impact on trade of PTAs.

some papers take into account the unobservable heterogeneity of countries, others the effect of zero-trade flows, the endogeneity of PTA variable or the persistency of trade; however, no paper appears to consider all these problems together, and thus the resulting estimates must be considered with extreme caution.

The paper is organized as follows. Section 2 introduces the gravity specification in its basic form. Section 3 presents the results related to the trade impact of PTAs estimated in the various papers, grouping the studies in those considering PTAs involving European, North American, Asian and "other countries". Section 4 provides a critical evaluation of the choices made with respect to the approaches chosen to carry out the empirical analyses in the surveyed papers and, finally, section 5 concludes.

2. The gravity model

Gravity models of international trade were first developed by Tinbergen (1962) and Pöyhonen (1963). In its basic formulation, the gravity model explains bilateral trade flows in analogy to Isaac Newton's law of gravity, by the attraction of two countries' "masses" (measured by GDP and/or population), reduced by the "distance" (which is a proxy of transport costs) and expanded by preferential trade agreements between them and by other factors, as, for example, a common language or colonial ties.

Some recent studies (Bergstrand, 1989; Limao and Venables, 2001; Soloaga and Winters, 2001) contribute to the refinement of the traditional explanatory variables and to the addition of new ones. Others (Mátyás, 1997; Cheng and Wall, 2005; Breuss and Egger, 1999; Egger, 2000) improve the econometric specification of the model.

Some criticism for the lack of theoretical foundations has emerged. However, as empirical applications of the gravity model have grown, theoretical bases of the model have also been proposed. Indeed, Anderson (1979) derives a theoretical foundation for the gravity model based on constant elasticity of substitution (CES) preferences and goods that are differentiated by region of origin. Subsequent extensions (Bergstrand 1989; Deardoff, 1995) consider monopolistic competition or an Heckscher-Ohlin structure to explain specialization. Finally, Anderson and van Wincoop (2003, p.174) manipulate the "CES expenditure system to derive an operational gravity model with an elegantly simple form".

According to the generalized gravity model of trade, indicating with (i,j) a pair of countries, the volume of exports of country *i* towards country *j* (X_{ij}) depends on their incomes, measured by GDPs (Y_i and Y_j), their populations (N_i and N_j), their geographical

distance (D_{ij}), *h* preferential trade variables (P_{ij}), and a k-dimension vector of country-pair specific factors that affect trade (F_{ij}).

Despite the fact that the gravity model is formulated in the multiplicative form, most studies have estimated the gravity parameters using the following log-linearised form:

$$\ln(X_{ij}) = \alpha_0 + \beta_1 \ln(Y_i) + \beta_2 \ln(Y_j) + \beta_3 \ln(N_i) + \beta_4 \ln(N_j) + \beta_5 \ln(D_{ij}) + \sum_h \delta_h P_{ijh} + \sum_k \lambda_k F_{ijk} + u_{ij} \quad [1]$$

where u_{ij} is the error term.

The coefficient β_1 is expected to be positive, since a high level of the domestic output indicates a high level of availability of goods to be exported. Also β_2 is expected to be positive, because a high level of the importing country's income implies potentially high imports. The distance coefficient is expected to be negative, because distance is a proxy of transportation costs. The signs expected for populations are ambiguous; there is no empirical evidence of a consistent sign for β_3 and β_4 (Cheng and Wall, 2005). Indeed, in most papers β_3 and β_4 are expected to be positive because it is believed that larger countries trade more. However, it has been shown (Oguledo and Macphee, 1994) that if an exporter is large in terms of population it may either need its production to satisfy domestic demand, so that it exports less, or it may export more than a small country, as it is the case when large firms achieve economies of scale. The same reasoning can be applied to the case of the importing country: if it is large, it may either import less because it is likely that the domestic sector develops and makes the country self-sufficient, or it may import more because it cannot satisfy all domestic demand with its own production (Pusterla, 2007).

The *h* PTAs are generally represented in the model by adding dummy variables.

Finally, the typical country-pair specific factors that affect trade considered in gravity models are common border, language, religion and past colonial ties.[2]

3. The impact of PTAs on bilateral trade

In this section we review the literature analysing the trade impact of PTAs. As mentioned in the previous section, the main objective of most papers is to evaluate the trade creation effect of PTAs. Generally, this objective has been pursued by augmenting the gravity equation by a

[2] These determinants are made available by Andrew Rose on his website (http://faculty.haas.berkeley.edu/arose/RecRes.htm#Software) for more than 180 countries and over the period 1948-2000. The database provided by Rose contains two binary variable set equal to unity if the trade partners share a common language or border, two categorical variables indicating the number of islands and landlocked countries in the pair (assuming the value of 0, 1 and 2), and a set of dummies concerning colonial ties.

dummy variable equal to one if both countries in the pair are members of a specific PTA and zero otherwise.

Besides the assessment of the trade creation effect, many studies evaluate if, and to what extent, PTAs divert trade; this is done by adding two dummy variables: the first is equal to one if only the importer belongs to the specific PTA considered and zero otherwise, the second is equal to one if only the exporter belongs to the specific PTA considered and zero otherwise. These two dummies allow to assess the impact on the two countries' imports and exports, respectively, of the trade diversion effect due to a specific PTA.

Furthermore, it is worth noticing that the papers reviewed in this survey greatly differ in their aims and in the specifications of the gravity equation.

Regarding the aims pursued, gravity models are used for addressing various economic issues. Some of the studies aim mainly at testing the impact of specific PTAs.[3]

However, there are also many other studies in which PTAs are considered only as controlling variables. In particular, a number of these studies attempts to assess the impact on trade of currency unions,[4] to compare different econometric specifications of the gravity model in order to improve estimates,[5] to show that rules of origin are important barriers to trade (Augier et al., 2005), and that alternative conditions of entry predict subtle differences in key parameters, depending on whether goods are homogeneous or differentiated and whether there are barriers to entry (Feenstra et al., 2001), or to investigate the role of distance in the gravity regression (Marquez-Ramos et al., 2006).[6]

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[3] In particular, Endoh (1999) and Bayoumi and Eichengreen (1995) focus on the effect of the EU integration process on trade; Cipollina and Salvatici (2007), Persson and Wilhelmsson (2005), Nilsson (2002) and Verdeja (2006) investigate the impact of preferences granted by the EU; Adam et al. (2003) analyse the role of PTAs among Eastern European countries; Wall (2003), Gould (1998) and Jayasinghe and Sarker (2004) study the effect of NAFTA, while Lederman and Özden (2004) examine the effectiveness of preferences granted by the US. Other papers aiming at assessing the trade impact of PTAs in general are Bun and Klaassen (2002a), Koo et al. (2006), Brada and Mendez (1985), Rose (2004a), Skripnitchenko et al. (2004), Carrère (2006), Baier and Bergstrand (2005), Álvarez-Coque and Selva (2006), Emlinger et al. (2006), Sandberg et al. (2006), Faruqee (2004), Martinez-Zarzoso and Nowak-Lehmann (2003), De Santis and Vicarelli (2006), Martinez-Zarzoso (2003), Frankel et al. (1995), Clarete et al. (2003), Lee et al. (2005), Martinez-Zarzoso et al. (2006), Pusterla (2007), Soloaga and Winters (2001), Tang (2005), Lee and Park (2005), Cernat (2001), Elliott and Ikemoto (2004), Martinez-Zarzoso (2003), Amurgo-Pacheco (2006), Kurihara (2003), Oguledo and MacPhee (1994), Kirkpatrick and Watanabe (2005), Carrillo and Li (2002).

[4] Saiki (2005), Bun and Klaassen (2006), Yeyati (2003), Rose and van Wincoop (2001), Rose (2000), Micco et al. (2003), Melitz (2001), Rose and Engel (2002), Pakko and Well (2001), Helpman et al. (2007), Katayama and Melatos (2006), Frankel and Rose (2000), Nitsch (2002).

[5] Cheng and Wall (2005), Breuss and Egger (1999), Sanso et al. (1993), Siliverstovs and Shumacher (2006), Bergstrand (1989), Egger (2005), Westerlund and Wilhelmsson (2006), Santos Silva and Tenreyro (2005), Tenreyro and Barro (2003), Baier and Bergstrand (2002), De Benedictis and Vicarelli (2004), De Sousa and Disdier (2002).

[6] Other studies have more specific focuses, such as, for example, the examination of the feature and the transition of trade relations in the Asia-Pacific region during the post-World War II period (Endoh, 2000), the assessment of the impact on trade of disintegration of three federations of Central and Eastern European countries (Fidmurc

Regarding the specification of the gravity equation, due to the lack of a strong and univocal theoretical foundation, the variables used differ from study to study.

As far as the dependent variable is concerned, some authors use total (or average) trade while others adopt data on exports or imports only. [7]

With regards to independent variables, the most used variables are GDPs[8] as proxy of countries' income. GDPs of the importer and the exporter enter into the regression separately, but there are authors considering a single variable, given by the product of the GDPs of the two countries (Rose, 2004a; Faruqee, 2004).

There is no agreement about the second measure of "masses", measured either by populations or by GDPs per capita. Moreover, some studies consider separately two variables, one for the importer and the other for the exporter, while others use the product of them, as for the GDPs.

Furthermore, the distance[9] among the capital cities of the two trading partners is used to proxy transportation costs, even if it could be subject to criticism on the grounds, for example, that transport by land and sea is assumed to have the same cost.

The choice of the country-pair specific factors affecting bilateral trade varies widely. Dummies for trading partners that share a common language and a land border are almost always considered. Often binary or categorical variables for trading countries that are islands or landlocked are also considered, as well as dummy variables indicating colonial ties.

Moreover, some authors consider a "remoteness" variable, that is a proxy of multilateral trade resistance (in addition to country-pair specific ones),[10] given by the

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and Fidmurc, 2003), the analysis of the debt renegotiation on international trade (Rose, 2005), the investigation of the integration within the EU zone (Sissoko, 2004), the impact of the EU enlargement (De Benedictis et al., 2005; Papazoglou et al., 2006), the effect of Mexico entry into NAFTA (Krueger, 1999), the analysis of the determinants of Bangladesh (Rahman, 2005), Asian (Kien and Hashimoto, 2005), Iceland (Kristjàndòttir, 2005), Czech Republic, Hungary and Poland (Jakab et al., 2001), and South Korea trade (Sohn, 2005), the role of innovation on trade (Eaton and Kortum, 1997), the trade relation between the EU and the Asian countries (Ronchi and Filippini, 1999).

[7] The main criticism brought against the use of imports as the dependent variable is that, since imports are recorded using c.i.f. prices (including transport and insurance costs), the variable measuring transport costs (distance, in most case) is correlated with the error term, thus generating a problem of inconsistency; conversely, as long as exports are calculated f.o.b., they do not cause any consistency problem (Pusterla, 2007).

[8] Nilsson (2002 and 2005) considers GNP and GNP per capita rather than GDP and GDP per capita, respectively.

[9] The distance variable is often computed considering the "great circle" measure (Rose, 2000; Rose, 2005; Siliverstovs and Schumacher, 2006; Melitz, 2001; Amurgo-Pacheco, 2006; Pusterla, 2007; Subramanian and Wei, 2005; Carrillo and Li, 2002; Agostino et al., 2007).

[10] However, according to Anderson and van Wincoop (2003) the remoteness variable, intended to reflect the average distance of a country from all trading partners, is "disconnected" from the theory because it does not account for border barriers to trade.

weighted sum of the distance between one country and the others, where the weights are the GDPs of the other countries.[11]

A variable given by the country's area is sometimes included[12] as well as the stock of infrastructures in the two trading partners.[13]

Furthermore, the (absolute or squared) difference in per capita GDP of the two trading partners and an index of similarity between the two GDPs[14,15] are sometimes considered in order to test the Linder hypothesis according to which countries with similar demand patterns are more likely to trade with each other.

Dummy variables for the pairs that have a common currency are also often used, especially when assessing the effect of using a common currency is the main aim of the paper.

Various authors include the weighted average of tariff rates in the gravity equation, by using import shares as weights (Fukao et al., 2002; Mayer and Zignano, 2005; Oguledo and MacPhee, 1994).[16]

A limited number of papers includes three other important variables: the price index of each country,[17] the exchange rate between the currencies of the two countries[18] and the exchange rate volatility.[19]

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[11] Carrère (2006), Sandberg et al. (2006), Fidmurc and Fidmurc (2003), Feenstra et al. (2006), Baier and Bergstrand (2002a and 2005), Ghosh and Yamarik (2004), Rauch and Trindade (1999), Martinez-Zarzoso et al. (2006), Melitz (2002), Rauch (1999), Pusterla (2007), Santos Silva and Tenreyro (2005), Soloaga and Winters (2001), Marquez-Ramos et al. (2006), Kimura and Lee (2006), Krueger (1999).

[12] Koo et al. (2006), Rose (2004a), Medved (2006), Ghosh and Yamarik (2004), Goldstein et al. (2003), Clarete et al. (2003), Lee et al. (2005), Melitz (2001 and 2002), Rose and Engel (2002), Saiki (2005), Subramanian and Wei (2005), Kurihara (2003), Lederman and Özden (2004), Lee and Park (2005), Kirkpatrick and Watanabe (2005), Baxter and Kouparitsas (2006).

[13] Acosta-Rojas et al. (2005), Carrère (2006), Martinez-Zarzoso and Nowak-Lehmann (2003), Sissoko (2004).

[14] Adam et al. (2003), Martinez-Zarzoso and Nowak-Lehmann (2003), Breuss and Egger (1999), Ghosh and Yamarik (2004), Tang (2005), Kirkpatrick and Watanabe (2005), Elliott and Ikemoto (2004), Carrillo and Li (2002), Medved (2006).

[15] In particular, Adam et al. (2003), De Benedictis et al. (2005), Breuss and Egger (1999), Medved (2006), Grünfeld and Moxnes (2003) consider the following similarity index between countries i and j at time t: $sim_{ijt} = \left[1 - \left(\frac{GDP_{it}}{GDP_{it} + GDP_{jt}}\right)^2 - \left(\frac{GDP_{jt}}{GDP_{it} + GDP_{jt}}\right)^2\right]$ where GDP_i and GDP_j are GDP of country i and country j, respectively, while Baxter and Kouparitsas (2006) consider the "uncentered correlation" index calculated employing industry n's share of country i's GDP s_{ni}, that is $sim_{ij} = \frac{\sum_{k=n}^{N} s_{ki} s_{kj}}{\sqrt{\sum_{k=n}^{N} s_{ki}^2} \sqrt{\sum_{k=n}^{N} s_{kj}^2}}$.

[16] Oguledo and MacPhee (1994) consider the average of the ad-valorem tariffs.

[17] Bergstrand (1985 and 1989), Gould (1998), Oguledo and MacPhee (1994), Fazio et al. (2005), Bayoumi and Eichengreen (1995).

[18] Bun and Klaassen (2002a), Carrère (2006), Martinez-Zarzoso and Nowak-Lehmann (2003), Bergstrand (1985 and 1989), Gould (1998), Micco et al. (2003), Rahman (2005), Saiki (2005), Sissoko (2004), Kien and Hashimoto (2005), Fazio et al. (2005), Jakab et al. (2001).

In a small number of contributions a dummy indicating GATT/WTO membership is also included,[20] to test if GATT/WTO membership increases/decreases trade.

Finally, other variables are also employed. For example, Ghosh and Yamarik (2004) consider the absolute difference in population density; Frankel et al. (1995) add the land-labour ratio; Siliverstovs and Shumacher (2006) and Lennon (2006) consider the average of years of schooling; Marquez-Ramos et al. (2006) include a technological indicator; Eaton and Kortum (1997) augment the gravity model by an indicator of research investments and human capital; Emlinger et al. (2006) consider a multinomial variable corresponding to the degree of perishability of the products.

Obviously, these differences in the model specification make a direct comparison of the results obtained in the various papers difficult.

3.1 PTAs involving European countries

The trade agreements promoted by European countries addressed in the contributions reviewed in this survey are the following: the European Economic Community (EC/EU); the European Free Trade Association (EFTA); the Council for Mutual Economic Assistance (CMEA/COMECON), the Central European Free Trade Agreement (CEFTA), the Baltic Free Trade Area (BAFTA), the agreement between the EU and the African Caribbean and Pacific (ACP) countries, the Euro-Mediterranean Agreements (EUROMED),[21]and, finally, the Generalised System of Preferences (GSP).

A large body of literature analyses the impact on trade of the creation and the enlargement of the EEC/EU. In general, the samples of countries considered consist of all

[19] Bun and Klaassen (2002a), Fidmurc and Fidmurc (2003), Baxter and Kouparitsas (2006), Yeyati (2003), Thom and Walsh (2002), Rose (2000), Nitsch (2002), Saiki (2005), Pakko and Wall (2001), Tang (2005), Kurihara (2003), Managi et al. (2005).

[20] Rose (2004a and 2004b), De Santis and Vicarelli (2006), Pusterla (2007), Subramanian and Wei (2005), Lederman and Özden (2004), Helpman et al. (2007).

[21] Signed in 1949, CMEA/COMECON was made up of the Soviet Union, Bulgaria, Czechoslovakia, Hungary, Poland, Romania, Albania, East Germany, Mongolia, Cuba and Vietnam. It was in force until 1991. CEFTA was signed by Poland, Hungary and Czech and Slovak Republics in 1992. Slovenia , Romania, Croatia and Republic of Macedonia joined later. In 2004 most of the countries (the exceptions are Croatia and Macedonia) joined the EU and thus left CEFTA. BAFTA is an agreement between Estonia, Lithuania and Latvia and it came into force in 1994. BAFTA ceased to exist when its members joined the EU in 2004. EEC and ACP countries signed their first agreements in 1969 during the Yaoundé Convention. In 1975, the Yaoundé agreements were replaced by the Lomé Convention, followed in 2000 by the Cotonou Partnership Agreement, which will regulate ACP-EU cooperation for the next 20 years. As regards EUROMED, Euro-Mediterranean Partnership (Barcelona Process) started in 1995. This partnership consists of 25 EU states and 10 southern Mediterranean states (Algeria, Egypt, Israel, Jordan, Lebanon, Morocco, Palestinian Territories, Syria, Tunisia, Turkey), with Libya having observer status since 1999. The Euro-Mediterranean Partnership comprises two complementary dimensions: a) a bilateral dimension: the EU carries out a number of activities bilaterally with each country; b) a regional dimension: regional dialogue represents one of the most innovative aspects of the Partnership, covering at the same time the political, economic and cultural fields (regional co-operation).

nations for which data are available (EU, USA, Canada, Asian countries, Central and South American nations) or in the EU members and their main trading partners.[22,23]

As regards data aggregation, many studies focus on total trade. [24,25]

The analyses cover different periods, even if most studies use long time spans (usually from the 60s to the 90s, or later),[26] while others limit the analyses to the 80s and the 90s.[27] Few papers consider the last two decades.[28]

With regards to the estimation methods, it is worth noting that a high number of papers presents cross-section results obtained from OLS or GLS estimators. This is done even when available data have both cross section and time dimensions.[29] There are few studies that use panel data, employing fixed effects, [30] random effects[31] and Hausman-Taylor (1981)[32] estimators. The fixed effects model takes into account the heterogeneity of each single

[22] Thom and Walsh (2002) and Kristjànsdòttir (2005) analyse only trade with major partners of Ireland and Iceland, respectively.

[23] The exceptions are the analysis of Bun and Klassen (2002a), Aitken (1973), De Sousa and Disdier (2002), De Benedictis et al. (2005) and Sapir (2001), who all limit their analyses to trade in Europe. In particular, De Sousa and Disdier (2002) consider imports of Hungary, Romania and Slovenia from EU and CEFTA countries over the period 1995-98; De Benedictis et al. (2005) consider 8 Central and Eastern European countries (CEECs) as reporting countries and EU15 and CEECs as trading partners over the period 1993-2003.

[24] The exceptions are Cipollina and Salvatici (2007), Koo et al. (2006) and Skripnitchenko et al. (2004), who limit the analysis to agricultural trade, Alvarez-Coque and Martì-Selva (2006) and Emlinger et al. (2006), who analyse fruit and vegetables trade, Eaton and Kortum (1997), Mayer and Zignano (2005) and Fitzsimons et al. (1999), who are interested in manufacturing trade, and Walsh (2006) who analyses trade in services. Furthermore, Mayer and Zignano (2005) also consider trade from north to north, north to south, south to north and south to south countries.

[25] Furthermore, Augier et al. (2005), Bergstrand (1989), Siliverstovs and Shumacher (2006) distinguish between manufacturing and agricultural trade, while Managi et al. (2005) conduct a study using data at 8-digit HS level. In more details, Siliverstovs and Shumacher (2006) consider separately all products combined, agriculture, mining and quarrying, and manufacturing products as a whole. Then they divide the latter by 25 3-digit ISIC Rev.2 industries. Bergstrand (1989) considers 1-digit SITC industry groups.

[26] Bun and Klassen (2002a and 2006); Endoh (1999); Carrère (2006); Persson and Wilhelmsson (2006); De Santis and Vicarielli (2006); Frankel et al. (1995); Thom and Walsh (2002); Sapir (2001); Lee and Park (2005); Gaulier et al. (2004); Bayoumi and Eichengreen (2005).

[27] Siliverstovs and Shumacher (2006); Martinez-Zarzoso et al. (2006); Marquez-Ramos (2006); Krueger (1999); Elliott and Ikemoto (2004); Cheng and Wall (2005); Martinez-Zarzoso and Nowak-Lehmann (2006); Carete et al. (2003); Rauch and Trindade (1999), Acosta-Rojas et al. (2005).

[28] Micco et al. (2003); Amurgo-Pacheco (2006); Pusterla (2007); Sissoko (2004); Managi et al. (2005); Kien and Hashimoto (2005); De Benedictis et al. (2005); Papazoglou (2006); Martinez-Zarzoso (2003); Westerlund and Wilhelmsson (2006); Mayer and Zignano (2005).

[29] Nilsson (2002), Eaton and Kortum (1997), Sapir (2001), Ghosh and Yamarik (2004), Sanso et al. (2003), Breuss and Egger (1999), Fidmurc and Fidmurc (2003), Elliott and Ikemoto (2004), Frankel et al. (1995), Clarete et al. (2003), Marquez-Ramos et al. (2006), Endoh (1999).

[30] Cheng and Wall (2005); Adam et al. (2003); Persson and Wilhelmsson (2006); De Benedictis et al. (2005); Martinez-Zarzoso and Nowak-Lehmann (2003); De Santis and Vicarelli (2006); Wall (2003); Westerlund and Wilhelmsson (2006); Walsh (2006); Martinez-Zarzoso (2006); Micco et al. (2003); Pusterla (2007); Lee and Park (2005); Kien and Hashimoto (2005); Gaulier et al. (2004); Fitzsimons et al. (1999), Jakab et al. (2001); Verdeja (2006); De Sousa and Disdier (2002).

[31] Adam et al. (2003); Papazoglou et al. (2006); Walsh (2006); Martinez-Zarzoso (2006); Lee and Park (2005); Kien and Hashimoto (2005); Fitzsimons et al. (1999); Jakab et al. (2001); De Sousa and Disdier (2002).

[32] Carrère (2006); Walsh (2006); Sissoko (2004); Kien and Hashimoto (2005); Verdeja (2006); De Sousa and Disdier (2002).

country and of each pair of countries. In fact, trade flows between two nations are likely to be affected by country and country-pair variables that are often not observable, such as the propensity to export (import) or the preference of a nation for another country products. If these specific effects are not properly taken into account, the resulting estimates could be biased. In the random effects model the specific effects are treated not as parameters to be estimated as in the fixed effects model, but as part of the error term. The most relevant difference between the fixed effects and the random effects models concerns the hypothesis of correlation between individual effects and explanatory variables; if the individual effects and the explanatory variables are not correlated, then fixed effects and random effects estimators are both consistent, but the fixed effects estimator is not efficient. If the individual effects and the explanatory variables are correlated, then only the fixed effects estimator is consistent. This hypothesis could be tested by the Hausman test, where the null hypothesis is the absence of correlation between individual effects and explanatory variables. The Hausman-Taylor (1981) estimator considers that only a subset of the independent variables are correlated with individual effects.

Other panel data estimators are also considered. Bun and Klaassen (2002a), De Benedictis et al. (2005), Martinez-Zarzoso (2006) and Micco et al. (2003) estimate a dynamic equation using the LSDV (Bun and Klaassen, 2002a), the GMM-system (De Benedictis et al., 2006; Mertinez-Zarzoso, 2006) and the Arellano-Bond (Micco et al., 2003) estimators. Bayoumi and Eichengreen (1995) estimate a first difference specification by OLS.

Moreover, a number of authors considers a nonlinear specification of the gravity equation, which is estimated by the Poisson Quasi Maximum Likelihood estimator (PQML)[33] (Westerlund and Wilhelmsson, 2006; Siliverstovs and Shumacher, 2006) or by the Nonlinear Least Square estimator (NLS) (Marquez-Ramos et al., 2006). Rauch and Trindade (1999), Amurgo-Pacheco (2006), Rauch (1999), and Soloaga and Winters (2001) use the log-linear specification of the gravity equation, and consider a Tobit model, [34] whereas Cipollina and Salvatici (2007), Gaulier et al. (2004) and Emlinger et al. (2006) employ a two-step Heckman

[33] The Poisson regression model is used for count data. The underlying assumption is that the probability that the dependent variable assumes a specific value conditional on independent variables follows a Poisson distribution. This model is estimated by the Maximum Likelihood estimator. However, often one of the hypotheses of the Poisson distribution, in particular the hypothesis that conditional variance is equal to conditional expectation, is violated. In this case, it is assumed that the Poisson distribution is not entirely correct, and the analysis is called Quasi Maximum Likelihood Estimation (Wooldridge, 2006).

[34] This model takes into account a kind of limited dependent variable that is continuous over strictly positive values but is zero for an important fraction of the population and is estimated by maximum likelihood estimator (Wooldridge, 2006).

(1978) procedure[35] to take into account zero trade flows. Furthermore, Eaton and Kortum (1997) consider 2SLS estimator for taking into account endogeneity of the wage variable that they include among regressors, while Thom and Walsh (2002) use an Error Correction Model estimator to take into account co-integration of trade and income variables.

Because of the heterogeneity in samples of countries, periods and estimation methods, the results are very different and can only be compared with caution.

3.1.1. EEC/EU and EFTA

As far as the EU/EEC is concerned, most studies find evidence of trade creation. In particular, it is mainly observed that the creation and the enlargement of the EU, *ceteris paribus*, fostered trade relationships not only between EU members but also between EU and non-EU countries.[36] However, there are also some studies that do not find evidence of an impact (positive or negative) on EU trade, since coefficients are not significant,[37] while others find a negative and significant effect.[38] When it is positive and significant, the impact of the creation and the enlargement of the EU is found to be smaller over the periods 1988-1996 (Martinez-Zarzoso and Newak-Lehmann, 2003) and 1992-2002 (Micco et al., 2003), the relative coefficient being equal to 0.04, while it is larger over the period 1999-2001 in the case of transport services (Walsh, 2006), the coefficient being equal to (a questionable value of) 14.9. If we focus on OLS estimates, we note that the EEC/EU coefficient ranges from 0.12 (Breuss and Egger, 1999) to 4.71 (Hassan, 2001), while using the fixed effects model it ranges from 0.04 (Martinez-Zarzoso and Newak-Lehmann, 2003; and Micco et al., 2003) to 1.84 (Adam et al., 2003). Even though comparisons must be made with prudence due to the

[35] This estimator takes into account the likely sample selection bias due to the fact that the process underlying the decision to export could be correlated with the model used to explain exports, that is the gravity model. If this correlation is verified, then estimates obtained disregarding this problem are biased. A method that allows us to take into account this issue is the two-step Heckman procedure (1978): in the first step the selection process is modelled by a probit model, then in the second step the gravity equation is estimated adding a correction factor, called Inverse Mills Ratio, retrieved from the probit estimates. Furthermore, Cipollina and Salvatici (2007) use a variant to the standard Heckman procedure proposed by Helpman et al. (2007), which simultaneously corrects for two types of potential biases: a sample selection bias and a bias from potential asymmetries in the trade flows between pairs of countries.

[36] Evidence of trade diversion for the EEC/EU agreement is found only by Westerlund and Wilhelmesson (2006), Kien and Hashimoto (2005) and Bayoumi and Eichengreen (1995). According to Westerlund and Wilhelmesson (2006) trade diversion coefficient probably captured only the effect on imports to Austria, Sweden and Finland and not changes in the imports of the EU.

[37] Koo et al. (2006); Skripnitchenko et al. (2004); Augièr et al. (2005); Ghosh and Yamarik (2004); Clarete et al. (2003); Eaton and Kortum (1997); Wall (2003); Westerlund and Wilhelmesson (2006); Raich and Trindade (1999); Walsh (2006); Krueger (1999); Hassan (2001); Bayoumi and Eichengreen (1995); Nitsch (2002).

[38] Thom and Walsh (2002), who studied Ireland's trade; Ronchi and Filippini (1999), for mineral fuels and chemical commodities; Rauch (1999) in the Tobit estimates; Soloaga and Winters (2001); Persson and Wilhelmsson (2006); Kien and Hashimoto (2005); Siliverstovs and Shumacher (2006), for wood, paper, photo and optical equipment industries.

differences in periods and countries sampled, from these results it seems that OLS tend to yield higher values for the impact of PTAs on trade. This is in line with the econometric theory, according to which in a panel data framework OLS overestimates parameters if individual effects and explanatory variables are correlated. Furthermore, EEC/EU coefficient ranges from 0.24 (Siliverstovs and Shumacher, 2006; manufacturing products) to 1.04 (Siliverstovs and Shumacher, 2006; agricultural products) if PQML is used, from 0.59 (Marquez-Ramos et al., 2006) obtained for 1995 to 1.80 (Marquez-Ramos et al., 2006) found for 1980 by using NLS, from 0.52 (Amurgo-Pacheco, 2006) when the effect on imports is estimated, to 1.18 (Amurgo-Pacheco, 2006) when the effect on exports is considered by estimating a Tobit model, and from 0.47 (primary goods) to 1.23 (consumption goods, Gaulier et al. 2004) by using a two step Heckman procedure.

As regards the EFTA dummy coefficient, it is mainly positive and significant, except in few cases where it is negative[39] or not significant.[40] According to Sanso et al. (1993), who find a significant EEC/EU and EFTA coefficient only up to 1972 and 1979, respectively, "both areas exhausted the advantages arising from their respective trade liberalization processes in those years and from that time membership is no longer a factor which clearly differentiates trade relations" (Sanso et al., 1993, p. 273). A similar opinion is expressed by Westerlund and Wilhelmsson (2006), who find a non significant EU coefficient over the period 1992-2002. They ascribe this evidence to the fact that the new members were already part of a free trade area with the EU before membership.

However, other authors (see notes 38 and 39) obtain a negative sign of the EU and the EFTA dummy coefficients and do not give any reasons to justify this questionable result.

Another strand of the literature considers the aggregate EU-EFTA PTA. In other words, the regression includes a dummy variable equal to one if the two countries in the pair belong to the EU or the EFTA. Bun and Klaassen (2002a and 2006), Gaulier et al. (2004) and Begstrand (1989) show that the impact of the aggregate EU-EFTA is positive, while in Jacab et al. (2001) and Nilsson (2002) it is negative. According to Bayoumi and Eichengreen (1995), the trade relationship between EEC and EFTA after a period of growth before 1959, falls in the five years following the establishment of EEC. "Since EFTA did not come into operation until 1965, it seems reasonable to attribute this contraction to the formation of

[39] Siliverstovs and Shumacher (2006), for agricultural trade; Rauch (1999), for commodities whose reference prices are quoted on organized exchanges (such as the London Meta Exchange); Lee and Park (2005).
[40] Rauch and Trindade (1999); Augier et al. (2005); Ghosh and Yamarik (2004); Sanso et al. (1993); Eaton and Kortum (1997); Sapir (2001); Soloaga and Winters (2001); Verdeja (2006); Nitsch (2002).

EEC" (Bayoumi and Eichengreen, 2001, p. 15). The estimated coefficient of this variable remains negative even after 1965, although it declines in absolute value.

3.1.2. Central and Eastern European PTAs

Various studies analyse Central and Eastern European PTAs. Endoh (1999), Gaulier et al. (2004) and Adam et al. (2003) consider the CMEA. Endoh (1999) estimates a year by year cross-section equation for 80 countries over the period 1960-1994, while Gaulier et al. (2004) and Adam et al. (2003) employ a fixed effects model over the period 1967-2001 and 1996-2000, respectively. Endoh (1999) and Gaulier et al. (2004) find that CMEA shows strong trade creation and diversion effects, since the coefficient of intra-bloc dummy is positive while coefficients of extra-bloc dummies (where only the exporter or the importer belongs to the agreement) are negative. Adam et al. (2003) find a negative CMEA coefficient over the period 1996-2000, which could be due to "the after-effects of the CMEA agreement" (Adam et al., 2003, p.11).[41]

Adam et al. (2003) and Augier et al. (2005) study the BAFTA and the CEFTA. Adam et al. (2003) consider a fixed effects estimation for the period 1996-2000 and find that BAFTA (whose coefficient is equal to 3.69) is more effective than CEFTA (1.52). According to the authors, these differences derive from the circumstance that BAFTA allows more liberalized trade between its members. Augier et al. (2005) estimate a gravity equation only for 1995 and 1999 and their results indicate that intra-CEFTA and intra-BAFTA trade effect was significantly higher across all products. CEFTA and BAFTA are also studied by Sissoko (2004) using the Hausman-Taylor model over the period 1988-2000. In this study, the CEFTA intra-bloc parameter is not significant, while the one for BAFTA is. Furthermore, for Baltic countries, when the estimated export trade diversion coefficients are significant they are negative and, hence, suggest a deterioration in competitiveness. For CEFTA trade diversion effect, Sissoko (2004) only finds significant results in 1995 and 1997 with a positive and negative impact, respectively. As regards the trade relationship between EU, CEFTA and BAFTA, Sissoko (2004) finds that the intra EU-BAFTA dummy is not significant over the period analyzed, whereas the intra EU-CEFTA one is significant and negative, although the values of the estimated parameters decrease over time. Moreover, a negative and significant sign of the CEFTA coefficient is found in Jakab et al. (2001) over the period 1990-1997.

[41] In fact, according to the Adam et al. (2003, p.11) "since the CMEA was a forced agreement, pre-1990 trade was inefficiently high, there was little competition, and the quality of the products was poor".

By estimating a gravity model for each year over the period 1990-1998, Fidmurc and Fidmurc (2003) find that trade between CEFTA countries strongly increased as a consequence of the introduction of the FTA. The same result is obtained by Pusterla (2007) who also provides evidence of trade diversion as a consequence of CEFTA, over the period 1988-2003. De Benedictis et al. (2005) find that intra-trade of Central and Eastern European Countries (CEECs) that joined the Union in 2004 raised during the period 1993-2003. A positive intra-bloc coefficient is also found by De Sousa and Disdier (2002), who analyse imports of Hungary, Romania and Slovenia from the EU and the CEECs.

3.1.3. Other European PTAs

A comprehensive analysis of PTAs involving the EU is provided by Persson and Wilhelmsson (2006). In particular, they focus on the effects of GSP, ACP and other agreements (Everything But Arms, South African FTA, drug regime, Euro-Mediterranean FTAs) [42] involving the European Union and estimate a fixed effects model over the period 1960-2002. Their results are mixed. In particular, they find that GSP has contributed to an increase in EU imports from developing countries (the coefficient is equal to 0.035 and significant). [43] On the other hand, a decrease of EU imports (-0.03) from the ACP countries follows the introduction of the Yaoundè Convention. However, if the aggregated effects are considered it can be concluded that all the countries, except those exporting to the EU under the drug regime, significantly benefit from trade preferences.

Nilsson (2002) analyses the effects of the GSP scheme granted by the EU and of the EU-ACP agreement over the period 1973-1992. He uses three year average trade data and finds that, for several EU countries, the impact of GSP preferences is highly significant, with positive or negative coefficients. In general, positive coefficients tend to occur at the beginning or at the end of the period analyzed. The only exceptions are the estimates regarding Belgium and the Netherlands, which are positive and significant in all periods, except 1984-86. With regards to the PTA in favour of ACPs, many EU countries show positive coefficients, which are either significant or insignificant. The impact is positive and significant for Belgium, the Netherlands, Spain and Portugal.

[42] The Everything But Arms initiative has been introduced in 2001 and gives tariff free and quota free access to all imports from the 49 Least Developed Countries. With regards to the South-African FTA, EU, Morocco, Israel and Tunisia signed agreements in late 1960s and early 1970s. These agreements were followed by the Cooperation Agreement signed with Maghreb (1976) and Mashrek (1977) countries. This agreement included non-reciprocal preferences and gave duty free access to many industrial and agricultural goods. The "drug regime" is a special arrangement with additional benefits for countries affected by the production and trafficking of illegal drugs, signed in 1991.

[43] A positive coefficient of GSP granted by EU countries is obtained also in De Santis and Vicarelli (2006).

Using a quantitative variable to measure preferential margins, Cipollina and Salvatici (2007) assess the impact of preferential trade agreements granted by EU to developing countries. Results show that GSP granted by EU, PTA for ACP countries and EU-Mediterranean PTAs significantly increase the probability of exporting to the EU market. Furthermore, the relatively larger increase in trade is found for GSP granted to Least Developed Countries (LDCs), followed by PTA for ACP countries and EU-Mediterranean agreements. As regards sectoral disaggregation, the larger effect of PTAs is observed for vegetable products and processed food industries.

Applying a Tobit model Amurgo-Pacheco (2006) estimates the effect of the trade agreements among European and Mediterranean countries. The dummy coefficient is always positive and significant. It is high for Syria and Turkey (3.090 and 4.017, respectively) when explaining the export flows, and for Syria (5.201) and Jordan (4.183) when the focus is on import flows. In the latter estimate, there is a negative and significant coefficient for Egypt. A positive impact of EU-Mediterranean agreements is also found in Pusterla (2007), Ålvarez-Coque and Martì-Selva (2006), and Gaulier et al. (2004). Pusterla (2007) finds evidence of trade creation even with third countries.

In their analysis on the access of Mediterranean countries to the EU market for fruit and vegetables, Emlinger et al. (2006) consider, rather than dummy variables, actual tariffs applied by the EU to its trading partners in order to better account for the preferences granted. The results they obtain show that the sensitivity of Algeria, Lebanon and Egypt to the preferential tariffs is very high, while Syrian, Tunisian and Jordan exports to the EU do not seem to be sensitive to tariffs, being the relative coefficients not significant.

Oguledo and Macphee (1994) obtain a negative coefficient for EU-Mediterranean and GSP granted by EU for 1976, while the coefficient of the Lomè variable is positive and highly significant. According to the authors, the Lomè Convention contains many other notable trade enhancing provisions besides preferential tariff reductions; in addition, many ACP countries have strong trading relations with European countries dating back to colonial times.

Using different estimation methods (OLS over two year periods, fixed effects, random effects), Verdeja (2006) estimates the impact of ACP, GSP and EU-Mediterranean PTAs over the period 1973-2000. Considering the cross-section results, the ACP coefficient is significant and positive in 8 out of 10 2-year periods, with coefficients ranging from 0.25 in 1999-2000 to 1.27 in 1993-95. Although the GSP coefficient is generally positive and significant, it is lower than that estimated for ACP countries. In particular, the GSP coefficient is significant in 3 out of 10 2-year periods and it ranges from 0.37 in 1981-83 to 0.75 in 1975-77. The EU-

Mediterranean coefficient does not reveal notable effects as it remains insignificant in 7 out of 10 periods (Verdeja, 2006).

When other estimation methods are used, there are no relevant differences in the estimated coefficients of PTAs. The country-pair fixed effects model yields again a positive and significant impact for ACP,[44] whereas every other agreement shows a worse performance- the EU-Mediterranean being insignificant and GSP negative and significant. According to Verdeja (2006), the GSP variable coefficient is negative because of the low utilization scheme by the GSP beneficiary countries, due to stringent rules of origin and administrative complications that make it very difficult for exporters to comply with the scheme's requirements.

3.2. PTAs involving North-American countries

The main regional trade agreement among North-American countries is the North America Free Trade Agreements (NAFTA), signed by Canada, United States (US) and Mexico in 1992. Other important FTAs are that between US and Israel, established in 1985, the African Growth and Opportunity Act (AGOA), the Caribbean Basin Initiative (CBI), the Andean Trade Promotion Act (ATP), [45] and the GSP.

Similarly to the studies that assess the EU's PTAs, the impact of PTAs established by the US has been analyzed by modelling the preferential treatment by a dummy variable that indicates the existence of a preference granted to the exporter by the importes. The only exception is Lederman and Özden (2004), who propose a preference utilization variable, computed as "the ratio of all exports entering under the program in that category to all exports from all eligible countries" (Lederman and Özden, 2004, p.11).

Most studies that assess the impact of North-American agreements on trade consider a large number of countries, except Fukao et al. (2002), who focus only on US trade flows, and Gould (1998), who limits his analysis to NAFTA countries only. The time span is usually very long (from the 60s or the 70s to the 90s).[46]

[44] Acosta-Rojas et al. (2005) also estimate the impact of ACP on trade reporting standardized coefficients, that measures the sensitivity of each independent variable to the dependent variable. The ACP coefficient is always positive, and ranges from 0.101 in 1994 to 0.160 in 1999.

[45] The AGOA was approved by the US Congress in 2000 to improve economic relations between the US and the economies of sub-Saharan African countries. The CBI was established in 1983 in order to provide several tariff and trade benefits to many Central American and Caribbean countries. The ATP was first signed in 1991 and provides duty-free and reduced-duty access to US markets for certain goods produced in Bolivia, Colombia, Ecuador and Peru.

[46] The exceptions are Koo et al. (2006) and Skripnitchenko et al. (2004) who estimate a gravity model only for 1999; Cheng and Wall (2005), Martinez-Zarzoso (2003), Clarete et al. (2003), Gould (1998), Martinez-Zarzoso

Most papers focuses on total trade.[47]

OLS estimation method and the fixed effects model are generally used, except Carrère (2006) and Kien and Hashimoto (2005) who use the Hausman-Taylor model, and Saiki (2005) who employs a random coefficient model. The latter estimation method allows the coefficient to vary across industries. This approach has two main advantages. First, it can alleviate the aggregation bias resulting from aggregating the industry level trade data to a country level. Second, the random coefficient model allows for parameter variation across industries, which is a more plausible assumption. In order to take into account persistency of trade flows, that is the fact that probably countries that traded a lot in the past trade a lot at present, Bun and Klaassen (2002a) and Martinez-Zarzoso et al. (2006) estimate a panel dynamic model by using the LSDV and the GMM-System estimators, respectively. Zero-trade flows are accounted for by Soloaga and Winters (2001) and Gaulièr et al. (2004), who consider a Tobit model and the Heckman two steps procedure, respectively. The Tobit model is also implemented by Lederman and Özden (2004) who also consider a two-step instrumental variable (IV) method to check for selection bias due to the preference dummy selection.[48] Tang (2005) uses the IV method to correct for endogeneity of exchange rate volatility. Finally, Siliverstovs and Schumacher (2006) and Marquez-Ramos et al. (2006) estimate the multiplicative gravity specification using the PMQL and the NLS estimator, respectively.

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et al. (2006), Saiki (2005), Marquez-Ramos et al. (2006), Jayasinghe and Sarker (2004) and Elliott and Ikemoto (2004) who consider the 80s and the 90s; Augier et al. (2005), Breuss and Egger (1999), Wall (2003), Fukao et al. (2002), Pusterla (2007), Tang (2005), Managi et al. (2005), Kien and Hashimoto (2005), Kristjànsdòttir (2005), Krueger (1999), Jakab et al. (2001), Cernat (2001), Mayer and Zignano (2005), Kien and Hashimoto (2005), who focus on the 90s; Siliverstovs and Schumacher (2006) who consider the 1988-1990 period; Soloaga and Winters (2001) who limit their sample to 1980-1986 years; Lederman and Özden (2004) who consider only 1997 and 2001; Hassan (2001) who consider only 1996 and 1997; Hilbun et al. (2006) who focus on 2001; Nouve and Staatz (2003) who use 1999-2002 data.

[47] The exceptions are: Koo et al. (2006), Skripnitchenko et al. (2004), Hilbun et al. (2006) and Nouve and Staatz (2003), who consider total agricultural trade, Mayer and Zignano (2005) who consider total trade in manufactured goods, Augier et al. (2005), who consider not only total trade but also agricultural trade and manufacturing trade. Furthermore, Fukao et al. (2002) use disaggregated manufacturing products (HS at 2-digit level), Siliverstovs and Schumacher (2006) consider all products, agriculture, mining and quarrying, manufacturing products as a whole and manufacturing products broken down into 25 3-digit ISIC Rev.2 industries, Gaulièr et al. (2004) use data disaggregated by primary, intermediate, consumption and capital goods, Saiki (2005) and Managi et al. (2005) consider also disaggregated commodities while Jayasinghe and Sarker (2004) consider only six selected products (red meat, grains, vegetables, fruit, sugar, oilseeds).

[48] This method consists of two steps. In the first step a probit model is estimated where the dependent variable is equal to one if there is a preferential trade agreement between the two countries in the pair and the regressors are: the same variables of the gravity specification, a dummy indicating if the exporting country maintains a formal alliance with the US, the average annual US total aid per capita to the country during the previous decade, a dummy for continental location. In the second step the gravity model is estimated using as IVs the retrieved probability obtained in the first step.

3.2.1. NAFTA

Many papers conclude that NAFTA has been effective in promoting trade between members.[49,50]

Comparing the EU creation and enlargement and the NAFTA, Breuss and Egger (1999) show that NAFTA has a greater impact on trade than the EU, the NAFTA coefficient being equal to 0.97 while the EU coefficient being equal to 0.12 only.

Managi et al. (2005) focusing on NAFTA effects, find that intra-bloc trade is increased in 47.8 and 27.1 percent of agricultural and non agricultural commodities, respectively. In contrast, the impact of EU on trade is significant in 43.5 and 7.1 percent of agricultural and non agricultural products, respectively. Thus, it seems that NAFTA is more effective in increasing intra-bloc export flows than the EU, and that the effects of regional economic integration is higher in agriculture than in non-agricultural sectors. The authors argue that these results could be due to the fact that "in agricultural trading, exporting countries use various types of trade restrictions much more than non-agricultural trading. These include export promotion programmes and importing countries have their own trade barriers to protect their domestic agricultural sector" (Managi et al., 2005, p. 8).

From an econometric point of view, it is interesting to note that these results are very sensitive to the estimation method used. For example, the estimated coefficient obtained by Jacab et al. (2001) with a Pooled Least Square regression is significantly negative, and becomes positive when they consider a fixed effects model. In the opinion of Jacab et al. (2001), this discrepancy is due to the heterogeneity bias of pooled regressions, which do not

[49] Only Ghosh and Yamarik (2004) and Siliverstovs and Shumacher (2006) obtain a negative coefficient.

[50] In more details, Martinez-Zarzoso (2003) finds that the NAFTA coefficient has been significant from 1965 onwards while, when limiting the analysis to Mexico, she finds that Mexican exports to the US and Canada countries are 161 percent higher than expected based on the basic gravity equation. According to Wall (2003), thanks to NAFTA, 29 percent more goods flowed from Canada to the United States and 14 percent more merchandise flowed from the United States to Canada. The results he obtains indicate also that NAFTA increased Canadian exports to Mexico by 12 percent and increased exports from Mexico to Canada by 48 percent. Furthermore, NAFTA led to a large decrease in Canada's exports to Europe and Asia, a decrease in Mexican imports from Europe, and a large increase in Mexican trade with Asia. To sum up, the only exception to the Vinerian trade creation and trade diversion effects is that the trade between Mexico and Asia increases as a consequence of NAFTA (Wall, 2003). In Jayasinghe and Sarker (2004) the NAFTA coefficient is high in the vegetables and red meat sectors. As for the trade of vegetables, the estimated coefficient rises to a significant value of 1.26 in 1998-2000. In red meat trade the coefficient of NAFTA is positive, significant during the entire period, and very high (1.78 during 1985-87 and 3.76 in 1998-2000). Hilbun et al. (2006) obtain an insignificant NAFTA coefficient. The authors ascribe this result to the fact that many agreements, such as NAFTA, "have specific time tables for the elimination of certain restrictions to trade. NAFTA had a 10-15 year goal of reducing/eliminating all external tariff barriers between trading members. As this research was conducted in 2006, and with NAFTA having been formed in 1994, the time limit has not yet been reached for total tariff elimination" (Hilbun et al., 2006, p.13-14).

control for the likely correlation between the unobservable individual effects and the explicative variables.

Evidence of trade diversion due to NAFTA is also found.[51]

3.2.2. Other North-American PTAs

Lederman and Özden (2004) merge in a single dummy NAFTA and Israel preferential trade agreements; moreover, they include a dummy for the GSP, CBI, ATP and AGOA agreements. They use dummy and preferential utilization variables (see above) to proxy trade preferences. The Tobit estimation results for 1997 show that all preference programs, except GSP, foster exports of the beneficiaries. A member of the FTA (NAFTA or US-Israel) exports almost three times more than a non member with identical characteristics. According to the authors, this effect might be rather large since the FTA dummy is weakly correlated with a border dummy (in fact, the border dummy is not significant). CBI beneficiaries export 136 percent more than other countries while the gain for Andean countries is 42%. On the other hand, GSP beneficiaries export 17 percent less. The Tobit results for 2001 show that the impact of AGOA is negative. In the opinion of Lederman and Özden (2004), this is due to the high negative correlation between distance and the AGOA dummy (distance between AGOA beneficiaries and US is high). The FTA (NAFTA and US-Israel) coefficient is almost identical to the one observed in 1997 while the CBI and ATP coefficients are higher. In the opinion of Lederman and Özden (2004) this is partially due to the expanded benefits of these programs in 2000, and to the increased capacity of exporters in taking advantage of trade preferences (Lederman and Özden, 2004).[52]

Nouve and Staatz (2003) evaluate the impact of AGOA on trade, using as PTA variable the AGOA dummy, which captures whether and when a country is declared AGOA

[51] Koo et al (2006), Skripnitchenko et al. (2004), Carrère (2006), Ghosh and Yamarik (2004), Clarete et al. (2003), Soloaga and Winters (2001), Krueger (1999), Elliott and Ikemoto (2004) and, for the exports, Kien and Hashimoto (2005). However, Martinez-Zarzoso et al. (2006) maintain that exports to non members do not change while imports from non members increase in the 90s. A positive extra-bloc NAFTA coefficient is found also by Pusterla (2007), Hilbun et al. (2006), Lee and Park (2004), Gaulier et al. (2004), Jayasinghe and Sarker (2004), Hassan (2001) for 1996, Cernat (2001) for 1994, and by Kien and Hashimoto (2005) for imports.

[52] Considering the Preference Utilization Ratios, the Tobit results for 1997 show that FTA and CBI have a positive and significant impact, whereas the ATP coefficient is not significant. In particular, an increase of 1% in utilization of preferences increases trade by almost 1% in FTA and by 1.2% for CBI members. The GSP coefficient remains negative. In 2001 the AGOA utilization variable is positive and significant, as well as the Andean, FTA and CBI coefficients, while GSP is still negative but lower in value. From the two-step-IV method, CBI, Andean and AGOA are positive and significant. The FTA impact is even larger than that for CBI, Andean and AGOA. From the Heckman model, utilization of FTA, CBI and AGOA has positive and significant effects on exports to the US, whereas GSP and Andean do not. Comparing the Heckman results with the Tobit ones, FTA, CBI and Andean utilization coefficients are significantly lower, whereas GSP and AGOA coefficients are higher, and GSP no longer has a negative effect. Other results indicate that trade between the United States and Israel is significantly higher across all products and years (Lederman and Özden, 2004).

eligible, and the VISA dummy, that indicates whether and when the eligible country has a visa system for apparel exports. The estimated AGOA dummy is not significant.[53] This can be due to the fact that AGOA was a relative recent initiative, and it may take longer before its impact can materialize in terms of increased agricultural exports. Furthermore, the implementation phase of AGOA coincides with a worldwide economic slowdown, and this may have lessened the real impact of AGOA on export performance in Sub-Saharian countries to the US (Nouve and Staatz, 2003). It is useful to point out that Nouve and Staatz (2003) exclude from their analysis textiles, which are, instead, eligible for AGOA treatment.

3.3. PTAs involving Asian countries

This section reviews the PTAs involving Asian countries. In particular, we consider the papers analyzing the impact of the Association of Southeast Asian Nations Free Trade Agreement (AFTA), the South Asian Association for Regional Cooperation PTA (SAPTA), the East Asia Economic Caucus free trade zone (EAEC), proposed in 1990 by Malaysia, the Economic Cooperation Organization (ECO), the Asia-Pacific Economic Cooperation (APEC), signed in 1989 and comprising of 21 members,[54] and the Gulf Cooperation Council (GCC), signed in 1981 and involving six Arab Gulf countries (Bahrain, Kuwait, Oman, Qatar, Saudi Arabia and the United Arab Emirates).

There are numerous studies that assess the impact of Asian PTAs, but few of them focus on the Asian countries only. These are Blomqvist (2004), who considers Singapore and its trading partners only, and Sohn (2005) who analyses Korea and its major trading partners only. In most papers, extended periods and total trade are considered, except for the few studies that focus on specific periods[55] or on a group of commodities.[56]

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[53] Similar to the AGOA dummy, the Visa coefficient is not significant (Nouve and Staatz, 2003).

[54] AFTA is an agreement by the Association of Southeast Asian Nations (Brunei, Cambodia, Indonesia, Laos, Malaysia, Myanmar, Philippines, Singapore, Thailand, Vietnam) signed in 1992. SAPTA has been signed in 1993 by SAARC member States (India, Pakistan, Nepal, Sri Lanka, Bangladesh, Bhutan and the Maldives) to promote and sustain mutual trade and economic cooperation among them. EAEC is expected to be composed of Asian economies: ASEAN, China, Hong Kong, Japan, Korea and Taiwan. The motive of this plan is to counteract the rapid institutionalization and trade liberalization of APEC led by the United States and also the strengthening of the regional political power of North America and Europe. ECO is an international organization involving ten Asian nations (Afghanistan, Azerbaijan, Iran, Kazakhstan, Kyrgyzstan, Pakistan, Tajikistan, Turkey, Turkmenistan and Uzbekistan) with the common purpose to establish a single market for goods and services. APEC includes most countries with a coastline on the Pacific Ocean (Australia, Brunei, Canada, Indonesia, Japan, Malaysia, New Zealand, Philippines, Singapore, South Korea, Thailand, United States, China, Hong Kong, Chinese Taipei, Mexico, Papua New Guinea, Chile, Peru, Russia, Vietnam) and aims at reducing trade tariffs to below five percent in the Asia-Pacific region, by 2010 for industrialized economies and by 2020 for developing economies.

[55] Koo et al. (2006) and Skripnitchenko et al. (2004) analyze the year 1999; Endoh (1999) and Sohn (2005) estimate the gravity equation for 1995 while Blomqvist (2004) for 1999; Siliverstovs and Schumacher (2006) consider the period 1988-1990; Mayer and Zignano (2005), Kien and Hashimoto (2005), Krueger (1999), Jakab

Estimation methods are mostly based on simple OLS, [57] except for some works where FE[58] or the Hausman-Taylor[59] panel estimators are considered. Moreover, there are papers taking into account zero trade flows with a Tobit estimation (Soloaga and Winters, 2001; Kirkpatrick and Watanabe, 2005) or with a two-step Heckman procedure (Gaulier et al., 2004) or by estimating a multiplicative form of the gravity equation by PQML (Siliverstovs and Shumacher, 2006).

The main conclusion that emerges from the literature on AFTA is that it enhances intra-bloc trade.[60] Results obtained in the papers considered also show increasing trade between AFTA member and third countries,[61] except in some studies where there is evidence of trade diversion (Clarete et al., 2003; Gaulier et al., 2004; Carrère, 2006; Hassan, 2001, for the 1996). Carrère (2006), in particular, obtains a negative coefficient for the dummy of the import trade diversion and a positive coefficient for the intra-bloc trade dummy and for the export trade diversion one.

Similar results are obtained for SAPTA.

With regards to EAEC, Endoh (2000) and Kirkpatrick and Watanabe (2005) obtain a positive coefficient. According to Endoh (2000), a positive and statistical significance of EAEC means that EAEC, a sub-region of APEC, has stronger trade relations than APEC.[62]

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et al. (2001), Cernat (2001) and Tang (2005) focus on the 90s while Soloaga and Winters (2001) on the last decade.

[56] Koo et al. (2006) and Skripnitchenko et al. (2004) consider only total agricultural trade; Siliverstovs and Schumacher (2006) disaggregate the sample in agricultural, mining and quarrying, and manufactured commodities; Mayer and Zignano (2005) use total manufacturing trade; Gaulier et al. (2004) divide the sample in primary, intermediate, consumptions and capital goods.

[57] Tang (2005) considers an instrumental variable model to take into account the likely endogeneity of exchange rate volatility, instrumented by the product, the sum and the average of inflation rates of both countries in the pair.

[58] Rahman (2005), Pusterla (2007), Lee and Park (2005), Kien and Hashimoto (2005), Gaulier et al. (2004), Jakab et al. (2001). Moreover, a random effects model is considered by Lee and Park (2005), Kien and Hashimoto (2005), Jakab et al. (2001).

[59] Carrère (2006) and Kien and Hashimoto (2005).

[60] Only in some studies the coefficient is not significant (Clarete et al., 2003; Blomqvist, 2004; Endoh, 2004; Rahman, 2005; Soloaga and Winters, 2001; Lee and Park, 2005; Jakab et al., 2001; Hassan, 2001, for the 1997) or negative (Kirkpatrick and Watanabe, 2005; Hassan, 2001, for the 1996; Pusterla, 2007, in the fixed effects estimation).

[61] Koo et al. (2006) find a positive trade diversion effect dummy coefficient. According to the authors, "the overall trade-diverting effect was positive, indicating that RPTAs do not displace agricultural trade with non-member countries. One reason may be because of a low degree of substitutability between traded products. Another potential reason for this result is that the trade-creating effect of RPTAs could increase overall demand to such an extent that the income effect outweighs the trade-diversion effect of the agreements" (Koo et al., 2006, p.415).

[62] According to Endoh (2000, p. 579) "member economies of APEC form a close trade relation area in the Asia-Pacific region, and EAEC forms an even closer trade relation area within APEC. The relation of these two areas may be illustrated as "a fried egg diagram" where APEC looks like the whole egg and EAEC is like the yellow part".

APEC is found to increase intra-bloc and extra-bloc trade,[63] suggesting that it is achieving its goals of open regionalism and augmenting total trade (Clarete et al., 2003). However, according to Sohn (2005), these results must be taken with caution because APEC is not a FTA, even though it is a "natural trading bloc" (Sohn, 2005, p.425); rather, "the significant and positive coefficient means that there exists a larger intra-APEC trade flow, which comes primarily from private business activities in the extended intra-regional production and/or from distribution networks, independent from any government effort of institutionalizing the integration" (Sohn, 2005, p.424-425).

Soloaga and Winters (2001) study the impact of GCC, and show a significant and positive coefficient for the intra-bloc dummy only in 1980 and in the period 1992-1996, and a negative coefficient for the extra-bloc export dummy. Different results are obtained by Kirkpatrick and Watanabe (2005), where the GCC intra-bloc coefficient is not significant but the GCC extra-bloc dummy coefficient is negative in 1970-71 and 2000-01 for imports and positive in 1996-97 and 2000-01 for exports.

Finally, the results obtained for ECO suggest that this PTA has a positive and statistically significant effect on intra-bloc trade in the early 1980s (Clarete et al., 2003).

3.4. PTAs among other countries

This section reviews the papers dealing with the impact of PTAs established by countries other than the EU, North American and Asian countries. In particular, the PTAs included are those among Central and South American countries, South Pacific countries and African countries.

PTAs involving Central and South American countries are the Andean Community of Nations (CAN), the Mercosur, the Caribbean Community and Common Market (CARICOM), the Central American Common market (CACM) and the Latin American Integration Association (LAIA). [64]

[63] Exceptions to this result are found by Siliverstovs and Shumacher (2006) who obtain a negative APEC coefficient for pottery industry in the OLS estimate, and for leather and footwear industries in the PMQL estimate, and by Pusterla (2007) who obtains a negative coefficient in the fixed effects estimation.

[64] CAN is a South American trading bloc comprising Bolivia, Colombia, Ecuador, Peru and Venezuela, founded in 1969. Mercosur is a RTA between Brazil, Argentina, Uruguay, Venezuela and Paraguay, founded in 1991. CARICOM is a Free Trade Association that came into effect in 1973, replacing the 1965-1972 Caribbean Free Trade Association (CARIFTA). It is composed by 15 members (Antigua and Barbuda, The Bahamas, Barbados, Belize, Dominica, Grenada, Guyana, Haiti, Jamaica, Montserrat, Saint Kitts and Nevis, Saint Lucia, Saint Vincent and the Grenadines, Suriname, Trinidad and Tobago). There are also five associate members (British Virgin Islands, Turks and Caicos Islands, Anguilla, Cayman Islands, Bermuda). CACM is a custom union between five nations of Central America (Guatemala, El Salvador, Honduras, Nicaragua and Costa Rica) established in 1960. LAIA is a Latin American trade integration association signed in 1980. Its main objective is

With regards to South Pacific countries, the PTAs considered are the Pacific Regional Trade Agreement (PARTA), the South Pacific Regional Trade and Economic Cooperation Agreement (SPARTECA), and the Closer Economic Relations (CER). [65]

The PTAs among African countries reviewed are the Southern African Customs Union (SACU), the West African Economic and Monetary Union (UEMOA), the Southern African Development Community (SADC), the East African Community (EAC), the Common Market for Eastern and Southern Africa (COMESA), and the Economic Community of West African States (ECOWAS).[66]

All the papers examined in this section consider numerous countries. The exception is Carrillo and Li (2002), that focus only on Andean and Mercosur countries. The analyses cover different periods of time.[67] As regards data aggregation, it emerges that most studies focus on total trade.[68] The most used estimation method is the OLS.[69]

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the establishment of a common market, in pursuit of the economic and social development of the region. Its members are Argentina, Bolivia, Brazil, Cuba, Chile, Colombia, Ecuador, Mexico, Paraguay, Peru, Uruguay and Venezuela.

[65] PARTA is an organization which aims at enhancing cooperation between the independent countries of the Pacific Ocean. It was founded in 1971. Member states are: Australia, the Cook Islands, the Federated States of Micronesia, Fiji, Kiribati, the Marshall Islands, Nauru, New Zealand, Niue, Palau, Papua New Guinea, Samoa, the Solomon Islands, Tonga, Tuvalu, and Vanuatu. Since 2006, New Caledonia and French Polynesia have been associate members. SPARTECA was signed in 1981 between Australia, New Zealand and countries of the South Pacific Forum. It allows duty free access for the products of Forum Island countries to Australia and New Zealand. CER is a free trade agreement between the governments of New Zealand and Australia. It is also known as the Australia New Zealand Closer Economic Relations Trade Agreement (ANZCERTA). It came into force in 1983 replacing the New Zealand Australia Free Trade Agreement signed in 1965.

[66] SACU was established in 1910 as a Customs Union Agreement between the then Union of South Africa and the High Commission Territories of Bechuanaland, Basutoland and Swaziland. With the advent of independence for these territories, the agreement was updated on December 11, 1969. SACU currently has five members (South Africa, Botswana, Lesotho, Swaziland and Namibia). UEMOA was created in 1994. It is a custom union and a monetary union among Benin, Burkina Faso, Côte d'Ivoire, Mali, Niger, Senegal, Togo and Guinea-Bissau. SADC is an inter-governmental organization that aims to build socio-economic cooperation and integration, as well as political and security cooperation, among 14 southern African countries (Angola, Botswana, Lesotho, Malawi, Mozambique, Swaziland, Tanzania, Zambia, Zimbabwe, Namibia, South Africa, Democratic Republic of Congo, Mauritius, Madagascar). EAC is a customs union in East Africa, consisting of Kenya, Uganda and Tanzania; Burundi and Rwanda will join on 1 July 2007. COMESA is a preferential trading area with twenty member states, from Libya to Zimbabwe, formed in December 1994 to replace a Preferential Trade Area which had existed since 1981. Nine of the member states formed a free trade area in 2000, with Rwanda and Burundi joining the FTA in 2004 and the Comoros and Libya in 2006. ECOWAS is a regional group, initially of sixteen countries (Benin, Burkina Faso, Cape Verde, Côte d'Ivoire, The Gambia, Ghana, Guinea, Guinea Bissau, Liberia, Mali, Niger, Nigeria, Senegal, Sierra Leone, and Togo), founded in 1975 in order to promote economic integration.

[67] Carrère (2006), Ghosh and Yamarik (2004), Frankel et al. (1995), Lee and Park (2005), Endoh, (1999), Kirkpatrock and Watanabe (2005), Nitsch (2002) consider a period extending from the 60s or 70s to the 90s; Soloaga and Winters (2001) focus on the 80s period, Tang (2005), Kien and Hashimoto (2005), Krueger (1999), Jakab et al. (2001), Cernat (2001), Martinez-Zarzoso and Nowak-Lehmann (2003), Mayer and Zignano (2005) on the 90s, and Brada and Mendez (1985) on the 70s; Cheng and Wall (2005), Martinez-Zarzoso (2003), Clarete et al. (2003), Martinez-Zarzoso et al. (2006), Pusterla (2007), Marquez-Ramos et al. (2006), Sandeberg et al. (2006), Carrillo and Li (2002) consider the 80s and 90s; Koo et al. (2006), Skripnitchenko et al. (2004), Gaulier et al. (2004) limit their analysis to 1999 while Hilbun et al. (2006) focus on 2001.

[68] The exceptions are: Koo et al. (2006), Hilbun et al. (2006) and Skripnitchenko et al. (2004), who analyse total agricultural trade, Mayer and Zignano (2005) who consider manufactured trade; Gaulier et al. (2004) who

3.4.1. Central and South American PTAs

A positive and significant intra-bloc trade effect is found for the CARICOM,[70] the Mercosur[71] and the CAN.[72]

In particular, Gaulier et al. (2004) find a very strong effect of Mercosur, corresponding to an increase in trade of more than 140%. Moreover, according to Clarete et al. (2003), who do not obtain evidence of import trade diversion effect, it seems that the increased intra-bloc trade within Mercosur has not reduced Mercosur imports from the rest of the world. Similarly, Lee and Park (2005), Kirkpatrick and Watanabe (2005) and Gaulier et al. (2006) obtain a positive and significant coefficient for the trade diversion effect dummy. From the results obtained by Cernat (2001), CAN intra-bloc dummy coefficient is more than 0.7, and the dummy coefficient of the export trade diversion is negative and ranges from -0.51 in 1998 to -0.25 in 1994. [73]

Coefficients for extra-bloc CARICOM trade is insignificant throughout the period analyzed, while intra-bloc trade coefficients are highly significant and positive.[74] Similarly, Sandberg et al. (2006) obtain a high value for CARICOM, equal to 4.9. The only exception to these results is provided by Lee and Park (2005) that find a negative and significant CARICOM coefficient.

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analyse primary, intermediate, consumption and capital goods, and, finally, Carrillo and Li (2002) who use the factor intensity classification (agriculture intensive, mineral intensive, labour intensive, capital intensive goods) proposed by the United Nations. Carrillo and Li (2002) also classify goods in homogeneous, reference priced and differentiated products, following Rauch (1999).

[69] The exceptions are Cheng and Wall (2005), Pusterla (2007), Lee and Park (2005), Kien and Hashimoto (2005), Jakab et al. (2001), Martinez-Zarzoso (2003, but for Mexico estimates only), Martinez-Zarzoso and Newak-Lehmann (2003) and Gaulier et al. (2004), that consider fixed effects estimation, Carrère (2006) and Kien and Hashimoto (2005), that employ the Hausman-Taylor model, Lee and Park (2005), Kien and Hashimoto (2005), Jakab et al. (2001) that consider RE. Furthermore, Martinez-Zarzoso et al. (2006) consider not only a panel methodology (FE and RE) but also a dynamic specification estimated by GMM-system, Soloaga and Winters (2001), Carrillo and Li (2002) and Kirkpatrick and Watanabe (2005) consider a Tobit model, Tang (2005) takes into account endogeneity of real exchange rate volatility using IV method, Marquez-Ramos et al. (2006) consider a multiplicative gravity specification estimated by Nonlinear Least Square, Gaulier et al. (2004) consider not only fixed effects estimation but also the Heckman two-step procedure to take into account zero-trade flows.

[70] Ghosh and Yamarik (2004), Martinez-Zarzoso (2003), Martinez-Zarzoso et al. (2006), Pusterla (2007), Marquez-Ramos et al. (2006), Cernat (2001), Sandberg et al. (2006), Nitsch (2002).

[71] Cheng and Wall (2005), Ghosh and Yamarik (2004), Frankel et al. (1995), Clarete et al. (2003), Pusterla (2007), Soloaga and Winters (2001), Lee and Park (2005), Marquez-Ramos et al. (2006); Kien and Hashimoto (2005), Gaulier et al. (2004), Jakab et al. (2001), Martinez-Zarzoso and Nowak-Lehmann (2003), Kirkpatrick and Watanabe (2005), Mayer and Zignano (2005).

[72] Carrère (2006), Frankel et al. (1995); Clarete et al. (2003), Pusterla (2007), Soloaga and Winters (2001), Lee and Park (2005), Marquez-Ramos et al. (2006), Gaulier et al. (2004), Cernat (2001), Nitsch (2002).

[73] Similar results in terms of sign is found for Mercosur. Over the period 1994-98 it appears that Mercosur doubled trade among members and reduced the extra-regional imports by one third.

[74] Their values suggest that trade between CARICOM countries was 10 times higher in 1994, and 85 time higher in 1998 than that between similar countries not part of that PTA.

Considering disaggregated data, results show that CAN has a significant effect only on capital intensive goods (Carrillo and Li, 2002).[75]

In general, CACM[76] and LAIA[77] are found to have a positive effect on trade. However, Endoh (1999) maintains that LAIA decreases trade among members and between members and non members. According to the Endoh (1999, p. 213), this negative effect is "a vivid reflection of the debt crisis that effected Latin American economies during the 1980s".

Finally, evidence of trade diversion is found for all the PTAs considered in this sub-section.[78]

3.4.2. South Pacific PTAs

The Australia-New Zealand CER is found to have a negative impact on trade by Cheng and Wall (2005), while a positive sign is obtained by Ghosh and Yamarik (2004), Pusterla (2007), Tang (2005) and Gaulier et al. (2004). In particular, Tang (2005) finds that the positive trade effect of the CER has gradually diminished since the mid-1990s. The CER agreement is also found to divert trade from non-members to members (Ghosh and Yamarik, 2004; Pusterla, 2007; Tang, 2005; Gaulier et al., 2004), even if Lee and Park (2005) obtain a positive trade diversion effect coefficient. SPARTECA seems to foster intra-bloc trade (Clarete et al., 2003)

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[75] In terms of reference priced products, CAN has a weak statistical significance in the aggregated category, while both CAN and Mercosur have a positive and significant effect on the trade of capital intensive goods. Furthermore, CAN has a negative and significant impact for mineral intensive products. There is no significant effect of preferential trade agreements in the case of homogeneous goods. When the CAN dummy coefficient is positive and significant it ranges from 0.43 for the aggregate differentiated category to 0.59 for reference priced capital intensive products, while the only significant coefficient for the Mercosur dummy is 0.81 (Carrillo and Li, 2002). According to Carrillo and Li (2002, p.16) "the results for the AC [Andean countries] seem to be consistent with the view that the new integration process has achieved an increase in the volume of trade of intra-industrial goods and particularly in the capital intensive ones. The lack of a statistically significant effect on agriculture, mineral and labour intensive products in each of the categories studied suggests that the capital intensive goods sub category has been the only one in which a statistically significant amount of trade creation has taken place. These may not seem so surprising since one would expect trade in agriculture and mineral intensive products to be driven mainly by others factors rather than integration, especially in countries with large natural resource endowments. Additionally, as mentioned in the previous section, the development of the capital intensive sector (and not the labour-intensive sector) has been one of the main objectives of the Andean integration process since its beginnings".

[76] Carrère (2006), Ghosh and Yamarik (2004), Martinez-Zarzoso (2003), Martinez-Zarzoso et al. (2006), Pusterla (2007), Soloaga and Winters (2001), Lee and Park (2005), Marquez-Ramos et al. (2006). In more details, the results found by Marquez-Ramos et al. (2006) show a decreasing effect of two Latin American agreements (CACM and CAN) on trade flows, and an increasing positive effect of CARICOM, Mercosur and NAFTA.

[77] Carrère (2006), Ghosh and Yamarik (2004), Soloaga and Winters (2001).

[78] That is, for the Andean Community (Koo et al., 2006; Hilbun et al., 2006; Skripnitchenko et al., 2004; Carrère, 2006; Ghosh and Yamarik, 2004; Clarete et al., 2003; Pusterla, 2007; Soloaga and Winters, 2001; Lee and Park, 2005; Gaulier et al., 2004; Cernat, 2001), the Mercosur (Carrère, 2006; Hilbun et al., 2006; Pusterla, 2007; Soloaga and Winters, 2001; Kien and Hashimoto, 2005; Krueger, 1999; Cernat, 2001), the CARICOM (Martinez-Zarzoso et al., 2006; Pusterla, 2007; Lee and Park, 2005), the CACM (Carrère, 2006; Hilbun et al., 2006; Ghosh and Yamarik, 2004; Martinez-Zarzoso et al., 2006; Pusterla, 2007; Lee and Park, 2005) and the LAIA (Carrère, 2006; Hilbun et al., 2006; Ghosh and Yamarik, 2004; Pusterla, 2007; Endoh, 1999).

and to reduce trade with non-members (Lee and Park, 2005). On the contrary, PARTA is found to foster trade with third countries (Lee and Park, 2005).

3.4.3. African PTAs

With regards to PTAs among African countries, Pusterla (2007) obtains a positive significant coefficient for the COMESA and UEMOA agreements. A positive COMESA coefficient is also found by Cernat (2001), who finds that imports by COMESA members from third countries were on average 30 percent higher than those predicted without accounting for the trade diversion dummy variable. Furthermore, Pusterla (2007) obtains evidence of trade diversion as a consequence of COMESA. Results also show that ECOWAS and SADC foster intra-bloc trade (Cernat, 2001),[79] and SADC increases trade with third countries (Cernat, 2001). Kirkpatrick and Watanabe (2005) obtain positive coefficients for the intra-bloc trade of African RTAs (1.08 for EAC, 1.49 for ECOWAS and 1.63 for SADC), suggesting that trade between countries belonging to an African RTA is higher than what would be expected, had they not be part of it. For EAC, the intra-trade coefficient is volatile over time, starting with 2.28 in 1970 and falling to 0.04 in 1978. This indicates that intra-EAC trade is more than expected until mid-1970s and drops in the following years (Kirkpatrick and Watanabe, 2005). Considering sub-periods, the EAC coefficient assumes the value of 2.22 in the 1970-71 period, sharply falls to 0.59 in 1978-79 and then increases again to 1.90 in 1996-97. According to Kirkpatrick and Watanabe (2005, p.155) "the changes of intra-trade coefficients for EAC coincide with the development of the East African integration". Finally, for ECOWAS the extra-bloc trade coefficient is negative in the 70s and positive in the 90s, while there is no evidence of trade diversion due to SADC (Kirkpatrick and Watanabe, 2005).

3.5. Assessments of the impact of PTAs in general

In several papers the goal of the analysis is not the assessment of the impact of the PTAs on trade *per se*, hence a PTA dummy is only included as a control variable. Generally, these papers consider an extensive sample of countries[80] over a long period.[81]

[79] A positive ECOWAS coefficient and significant until 1980 is found also by Nitsch (2002).
[80] The only exception is De Benedictis and Vicarelli (2004) who limit their analysis to exports from France, Germany, Italy and Spain.
[81] Except for Micco et al. (2003) who consider 1992-2002 years, Agostino et al. (2007) who focus on 1995-2003 years, Santos Silva and Tenreyro (2005) who analyse 1990, while Rose and Engel (2002) 1995, and Bayer and Bergstrand (2002) 1996, Grünfeld and Moxnes (2003) 1999. Furthermore, Freund and Weinhold (2004) limit their analysis to the years 1995 and 1999, Kimura and Lee (2006) use 1999-2000 data, while Lennon (2006) and Faruqee (3004) 1999-2002 data, Medved (2006) focuses on 2000-2002 years, Brada and Mendez (1985)

As regards disaggregated data, most papers focus on total trade.[82]

In general, the OLS estimation method is used, even if fixed effects method is also very frequent.[83] Furthermore, a dynamic gravity specification is estimated by using the Arellano-Bond estimator (Micco et al., 2003; Rose, 2004a) or by GMM-system (De Benedictis and Vicarelli, 2004). Rose (2004a) and Rose (2005) consider also the Tobit and Random effects estimator,[84] and Rose (2005) uses instrumental variables to take into account the hypothesis that trade and debt could be determined simultaneously.[85] The instrumental variable method is also used by Rose (2000) to take into account endogeneity of currency union and exchange rate volatility,[86] by Tenreyro and Barro (2003) to take into account the endogeneity of common currency variable.[87] Helpman et al. (2007) and Agostino et al. (2007) consider the two-step Heckman (1978) procedure (see note 35) to take into account zero-trade flows. [88] Furthermore, Santos Silva and Tenreyro (2005) estimate a non linear

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consider the 70s, Helpman et al. (2007) the 80s, and Mayer and Zignano (2005), De Benedictis and Vicarelli (2004), Egger (2005) focus on the 90s, whereas Nilsson (2005) considers the period 2001-2003.

[82] Agostino et al. (2007) consider also total agricultural trade and trade of ten 2-digit aggregation (SITC) of agricultural products, Feenstra et al. (2001) disaggregate the sample into differentiated, reference priced and homogeneous goods and Subramanian and Wei (2005) break down commodities in liberalized manufacturing (sectors with tariff higher than 5 percent in 1989 and zero in 2000 for both the US and the EU), protected manufacturing (sectors with tariff higher than 5 percent in both 1989 and 2000, and tariff lines decreased by less than 2), clothing, footwear and food and then estimate a Seemingly Unrelated Model in which each equation is relative to a commodity group. Furthermore, Lennon (2006) distinguishes between bilateral trade in goods and in services, while Grünfeld and Moxnes (2003) analyse trade in services only.

[83] Agostino et al. (2007), Melitz (2007), Micco et al. (2003), Helpman et al. (2007), Baier and Bergstrand (2005), Yeyati (2003), Rose (2004a), Rose (2004b), Rose (2005), Rose and can Wincoop (2001), Melitz (2002), Tenreyro and Barro (2003), Mansfield and Bronson (1997), Fazio et al. (2005), Katayama and Melatos (2006), De Benedictis and Vicarelli (2004), Goldstein et al. (2003), Medved (2003), Lee et al. (2005), Faruqee (2004), Baxter and Kouparitsas (2006), Pakko and Wall (2001), and Kurihara (2003) consider FE. Baier and Bergstrand (2002 and 2005) estimate also an average treatment effect by a two-step IV methods to take into account the endogeneity of the PTA dummy. In the first step a probit is considered and then, in the second step, the gravity equation is estimated by IV using predicted probabilities obtained in the first step. In the probit estimation independent variables are those of the gravity equation, political variables ("accountability indicator", that measures the extent to which citizen of a country are able to participate in the selection of its government; "political stability", that measures the ability of citizens to peacefully select and replace those in power; "government effectiveness", that measures the quality of bureaucracy and the credibility of government's commitments) and economic variables ("remoteness" of the country to the rest of the world; "difference in their capital-labour ratios"; "difference in the pair's capital-labour ratio from that of the rest of the world") (Baier and Bergstrand, 2005).

[84] The random effects model is also performed by Rose (2004b), Lee et al. (2005), and Subramanian and Wei (2005).

[85] The instrumental variables used are domestic and foreign CPI inflation rates, current accounts and government budget surplus/deficit (Rose, 2005).

[86] A first set of instrumental variables used is the product of the two inflation rates, their sum and the absolute value of their difference; then, Rose (2000) estimate a further regression using another set of instruments, consisting of the same variables in the previous set and also of the product, the sum and the absolute value of the difference between the two growth rates of M2.

[87] First they estimate the probability that a given country adopts the currency of a main anchor country. The instrumental variable is then obtained by computing the joint probability that two countries, independently, adopt the same currency (Tenreyro and Barro, 2003).

[88] The authors implement also the Wu-Hausman test to verified the endogeneity of PTAs variables, but the results show that the hypothesis of endogeneity is rejected in all estimations (Agostino et al., 2007).

specification of the gravity model by NLS and by PQML, and propose Tobit estimates for comparison.

The results show that an FTA is nearly always trade expanding. In fact, the relevant coefficient is almost always positive[89] and significant and ranges from 0.17 obtained by Freund and Weinhold (2004) in 1995, to 4.8 obtained by Brada and Mendez (1985) in 1976.

Rose (2004a) includes in his analysis a GSP dummy whose coefficient ranges from 0.04 (but not significant) to 0.86. However, in Rose (2004b) the GSP coefficient if found to be always negative. Similarly, Subramanian and Wei (2005) obtain a negative GSP coefficient as Rose (2004b) but a positive one when the WTO dummy (equal to one if the importer is a WTO member, but between the two pair countries there is no FTA or GSP agreement) is defined as mutually exclusive with FTA and GSP.[90] Goldstein et al. (2003) also estimate separate effects for GATT according to whether the importer or exporter is a member of GATT or a GSP preference giving or preference-receiving country. "Trade actually declines by 17% when the importer grants a preference to the exporting country and increases by 32% when the exporter grants a preference" (Goldstein et al., 2003, p.22).

Agostino et al. (2007) consider ordinary GSP dummy, a dummy equal to one if the exporting country benefits from more favourable conditions granted to Least Developed Countries (LDCs) within the GSP scheme, and a dummy for other non-reciprocal preferential trade agreements. From the Heckman two-step procedure results, the ordinary GSP coefficient is positive when considering total (agricultural and non agricultural) exports, but it becomes negative when splitting the sample by industries, except for the meat sector (it is positive and equal to 0.23). On the contrary, coefficient of GSP for LDCs when significant is positive, except for coffee, tea, cocoa and spices. Similarly, the other preference coefficient is always positive when significant, except for feeding stuff for the animals sector. Agostino et al. (2007) consider also a multinomial variable, which assumes higher values as the margin of the beneficial treatment increases. In more details, it is equal to: zero if the export flow from country i to country j receives no preferential treatment; one if it is regulated under the ordinary GSP only; two if there is a preference based on the GSP for LDCs only; three if the only preferential treatment received is that from other preferences; four if the trade flow may enjoy preferential treatment based on ordinary GSP and as well as under other preferences;

[89] A negative and significant FTA coefficient is found only by Grünfeld and Moxnes (2003) and by Lee et al. (2005).

[90] Furthermore, from their estimation by commodity groups it emerges that an FTA is positive and significant for liberalized and protected manufactured and clothing industry, whereas a GSP coefficient is positive (and significant) only for liberalized manufacturing and for the clothing industry, while it is negative for footwear and food industries (Subramanian and Wei, 2005).

five if the export flow is eligible for a preferential treatment under the GSP for LDCs and under other preferences. Results obtained by using Heckman two-step procedure show that passing from lower to higher levels of trade preference increases total exports, the relative coefficient being significant and equal to 0.27. Even when splitting the sample by industries, the multinomial variable coefficient if almost always significant, and when positive it ranges from 0.13, observed for fish, crustaceans, mollusc and other aquatic invertebrates, to 0.57, observed for sugar, sugar preparations and honey.

Nilsson (2005) compares the effect on LDC exports of EU and US preferential trade policies over the period 2001-2003. Results show that developing countries' exports gain more from EU preferential policies than from US ones. In particular, "the effects of EU trade policy are relatively larger for the poorest group of developing countries, which is dominated by LDCs that enjoy the most preferential access to the EU market through the EBA initiative" (Nilsson, 2005, p. 17).

Medved (2006) considers PTAs notified to WTO and not notified to WTO, bilateral agreements and multilateral agreements, and North-North, North-South and South-South agreements. He obtains a higher coefficient for PTAs notified (0.988) than for those not notified (0.667) to WTO, a similar coefficient for bilateral (0.846) and multilateral (0.870) agreements, a non significant coefficient for North-North agreements, and significant coefficients equal to 0.233 for North-South agreements and to 1.227 for South-South agreements (Medved, 2006).

Besides considering intra-bloc and extra-bloc dummies, Lee et al. (2005) introduce a "RTA-SingleOverlap" dummy "which is unity if both i and j belong to the same RTA, and either i or j exclusively belongs to another RTA with other countries" (Lee et al., 2005, p.33), and "RTA-GroupOverlap" dummy that is equal to one if both countries belong to multiple RTAs.[91] The results indicate that it is counterproductive to be a member of more than one RTA. Indeed, the estimated coefficient for "RTA-SingleOverlap" is negative and highly significant, indicating that "if a member forms another RTA, by creating overlapping RTAs, its additional trade with members of existing RTA(s) or with members of new RTA(s) is less than the additional trade formed between members belonging to a single RTA" (Lee et al., 2005, p.31). Similarly, the "RTA-GroupOverlap" is negative and highly significant,

[91]Although intra-bloc and extra-bloc RTA dummies are always considered, the other dummies are considered in different equations: RTA-singleover in a first estimate and RTA-multover in a second regression.

confirming that overlapping RTAs may not be as effective as a single RTA (Lee et al., 2005).[92]

Lennon (2006) estimates trade in services and trade in goods separately. Results indicate that participation in a RTA seems to be more important for trade in services than for trade in goods.[93]

Brada ad Mendez (1985) consider as PTA variable a categorical variable equal to 2 if both countries belong to the same preferencial agreement, and 1 if the two countries belong to different or no preference areas. Furthermore, they include a dummy (equal to 1 if both countries belong to the same preference agreements and 0 otherwise) multiplied by the GDP per capita of countries i and j and by the distance. The former coefficient "measures the effect of per capita income on the effectiveness of integration. If the coefficient is positive then the effect of integration on inter member trade increases with the level of development of the integrating countries, reflecting the higher proportion of tradable in their output" (Brada and Mendez, 1985, p.551-552). The latter coefficient measures "the effect of distance on the trade augmenting power of a customs union. The greater distance among members, the smaller, *ceteris paribus*, is the augmentation of their trade with each other" (Brada and Mendez, 1985, p. 552). Their results show a positive PTA coefficient (equal to 3.8 in 1970, 4.7 in 1973 ad 4.8 in 1976) indicating that integration reduces resistance to trade among member countries. The value of interaction between the PTA dummy and the product of GDPs per capita falls over time and in 1976 is not significantly different from zero. According to Brada and Mendez (1985, p. 552), this result could be due to "the worldwide increase in the prices of fuels and raw materials, since this increase then caused complementary trade in such goods among countries of different income levels to be weighted more heavily in total trade than competitive trade flows among developed countries". Finally, the value of interaction between the PTA dummy and the distance is always negative, indicating that the effects of

[92] Lee et al. (2005) consider also an "RTA-original" dummy equal to one if both countries are members of an RTA that expands its membership, "RTA-expansion" dummy equal to 1 if one of the two trading countries joins an existing RTA, and "RTA-duplicate" dummy equal to one if the two trading partners belong to different RTAs. Estimated coefficient of "RTA-original" is negative but insignificant, indicating that trade creation for original members is not significantly different from that for the intra-membership of RTA that never expands. "RTA-expansion" coefficient is negative, large (-0.635) and highly significant, indicating that trade creation for new members is much lower than that for the original members. Finally, "RTA-duplicate" coefficient is positive (0.166) and highly significant, indicating that trade is actually created more from the subset of the non members that belong to some other RTAs.

[93] Furthermore, when services are disaggregated by groups, the impact of PTAs on trade in goods and on trade in "other commercial services" are no longer significant when the GDP per capita variable is included, whereas the impact of the RTA on trade in "government services" is lower than that on goods (Lennon, 2006).

integration on trade reduce as distance between integrating countries increases (Brada and Mendez, 1985).

4. Empirical issues in the selected analyses

The results regarding the trade impact of PTAs discussed in the previous section are very heterogeneous, both in sign and magnitude. In addition, it is important to note that all these studies present at least one empirical problem.

First of all, except in Cipollina and Salvatici (2007) and Emlinger et al. (2006), the PTA variable is proxied by using the dummy approach. However, the approach of including PTAs in a gravity model by using dummy variables is problematic because the dummies capture a range of other country-pair specific effects contemporaneous with PTA implementation. Moreover, dummy variables treat all the countries that signed the preferential agreement as a homogenous group. This procedure does not take into account the heterogeneity due, for example, to the different rate of utilization of preferential margins. A further drawback of the dummy variable approach regards the fact that it does not discern among the different preferential trade policy *instruments* (preferential tariff margins, preferential quotas, reduced "entry prices") as well as it does not measure the *level* of trade preferences (i.e., dummies impose that the level of preferential schemes under GSP is the same of those under Euro-Mediterranean Agreements). Furthermore, PTAs may have different impacts on trade in different products. If total imports or exports are considered as dependent variables, this means to assume that trade flows in all commodities are affected in the same way by the set of independent variables used in the model. In addition, even if there is a preference for a specific product, for which the MFN duty is above zero, an exporter may choose to export the goods under the MFN tariff if, for example, the costs of fulfilling the "rules of origin" in order to benefit from the preferential tariff are too high (Medved, 2006). So far, these concerns are disregarded by almost all the papers which use dummies to proxy PTAs and consider aggregate data.

A strong effort is needed in this regard, to replace the dummy variable with variables providing more accurate information on the specific preference margin associated with the PTA, possibly considering differences on a product by product base. Lederman and Özden (2004) propose as proxy of the PTA a preference utilization variable, computed as the ratio of all exports entering under the program in that commodity category with respect to all exports of that category from all eligible countries. Even if this indicator allows to overcome

problems arising from the use of dummies, it is, however, a very "indirect" proxy of the PTA. For example, a good indicator of PTAs could be based on the difference between the (average) PTA tariff and the (average) MFN tariff. The main problem to be faced addressing this issue is the availability of data. However, recent databases have been developed to fill the gap. One of these datasets is MacMap-HS6 (Bouët et al., 2004), which provides an ad valorem equivalent measure of bilateral tariff duties and tariff quotas in 2001 and 2004 for 163 countries and 208 partners at 6-digit level of the Harmonized System (HS) classification. Furthermore, Bchir et al. (2005) compute ad valorem equivalent bound duties at 6-digit HS, for almost all WTO members. In addition, the "TRADEPREF" database (Gallezot, 2005b) associates EU imports in 2002 to the preferential regimes used. Moreover, the Integrated Tariff of the European Community (TARIC) database provides all information regarding tariff suspensions, tariff quotas, tariff preferences, preferential quotas, GSP applicable to developing countries, and other information concerning European trade regulations (Gallezot, 2005a). The latter dataset is used by Emlinger et al. (2006), who consider, as PTA variables, specific tariffs applied by the EU to each of its trading partners, while Cipollina and Salvatici (2007) compute the preferential margin variable using the MacMap dataset for 2004.

Secondly, the studies considered in this review address in an unsatisfactory way some econometric issues, the first of which is the country heterogeneity. This could be taken into account by using a fixed effects model, that allows to control for all factors that are fixed over time. Moreover, distance, which is meant to reflect the cost of trading between the countries of the pair, is recognised to be a poor measure of such costs. For example, in terms of trading costs per km, distance across land is not the same as distance across water. And furthermore, transport costs across less developed countries are not the same as across developed ones because of the differences in infrastructure stocks and quality (Pakko and Wall, 2001). A fixed effects model could be used to take into account such factors not properly measured by distance. Moreover, the fixed effects could be used to approximate the multilateral trade resistance index (De Benedictis and Vicarelli, 2004; Rose and van Wincoop, 2001) developed by Anderson and van Wincoop (2003). Formally, the gravity specification (1) for country pair (i,j) at time t becomes:

$$\ln(X_{ijt}) = \eta_{ij} + \beta_1 \ln(Y_{it}) + \beta_2 \ln(Y_{jt}) + \beta_3 \ln(N_{it}) + \beta_4 \ln(N_{jt}) + \beta_5 \ln(D_{ij}) + \sum_h \delta_h P_{ijth} + \sum_k \lambda_k F_{ijtk} + u_{ijt} \quad [2]$$

where η_{ij} indicates the country-pair fixed effects. Some authors consider country fixed effects α_i and γ_j rather than country-pair specific effects, that is they assume $\eta_{ij} = \alpha_i + \gamma_j$.[94] When individual effects are omitted, OLS estimates will be biased (in particular, they are overestimated) if individual effects are correlated with the regressors.[95] For example, results obtained by Egger (2005) show that OLS estimates exaggerates the importance of trading blocs. According to Egger (2005, p.887) "the OLS estimator erroneously attributes some effect to trading bloc membership, which in fact is due to other influences such as the true relative endowments with human or physical capital, which are difficult or even impossible to measure, or legal standards, which are likewise difficult to quantify. In the present dataset, the member countries in free-trade areas are among the richest in the OECD. On average, these countries also exhibit high legal standards, stable political systems, and according to available estimations they are well endowed with physical and human capital". According to the author, these issues are properly accounted for in both fixed effects and the Hausman-Taylor estimators. However, in the Hausman-Taylor model some independent variables are assumed to be correlated with the individual effects, while others are not. Since individual effects contain all factors that are specific to the trading partners, it would be unlikely that explanatory variables are not correlated with them. Hence, a fixed effects model could be less unreliable.

A fixed effects model has been often used in recent analyses; however these, on the other hand, disregard other econometric problems. One of these problems is the selection bias that can arise when two different "processes" are correlated. The first "process" is the selection equation, which describes the decision to export or not, while the second is the outcome "process" that generates the amount actually traded. In our case the outcome process is given by the gravity specification. Consider a generic framework of the equation of primary interest:

$$Y_1 = X_1'\beta_1 + \varepsilon_1 \qquad [3]$$

where Y_1 is only observed when Y_2=1. Suppose that Y_2 is driven by the following model:

$$Y_2^* = X_2'\delta_2 + \varepsilon_2 \qquad [4]$$

[94] For example, Yeyati (2003), Rose (2004b), Rose and van Wincoop (2001), Subramanian and Wei (2005), Tenreyro and Barro (2003), Lederman and Özden (2004), Fazio et al. (2005) consider country specific fixed effects, while Bun and Klaassen (2002a), Cheng and Wall (2005), Rose (2004a), Adam et al. (2003), Càrrere (2003), Persson and Wilhelmsson (2005), Faruqee (2004), Martinez-Zarzoso and Nowak-Lehmann (2003) consider country-pair fixed effects.

[95] One criticism to the fixed effects model is that it does not allow us to measure long run relationships. However, it can be argued that "the long run, the period over which no factors related to trade are fixed, is a length of time that is of no interest to policymakers or anyone else" (Pakko and Wall, 2001, p.41).

with $Y_2 = 1$ if $Y_2^* > 0$ and $Y_2 = 0$ if $Y_2^* \leq 0$.

If we consider the expected value for Y_1, we have:

$$E(Y_1 \mid Y_2 = 1) = X_1'\beta_1 + E(\varepsilon_i \mid Y_2 = 1) = X_1'\beta_1 + E(\varepsilon_i \mid Y_2^* > 0) = X_1'\beta_1 + E(\varepsilon_i \mid \varepsilon_2 > -X_2'\delta_2) \quad [5]$$

If ε_1 and ε_2 are correlated, then $E(\varepsilon_i \mid \varepsilon_2 > -X_2'\delta_2) \neq 0$. This implies that estimates obtained disregarding this correlation are biased. To overcome this problem, an Heckman (1978) estimator could be used. This approach has been carried out by Lederman and Özden (2004), Gaulier et al. (2004), Helpman et al. (2007), Emlinger et al. (2006), Cipollina and Salvatici (2007) and Agostino et al. (2007).

Anyway, even if this correlation does not exist, ignoring zero-trade flows could affect results. In more details, the general approach is that of log-linearizing gravity specification and treating zero-trade flows as missing values, that is sweeping them out from the analysis. However, this procedure could yield biased estimates. In this sense, a Tobit model may allow to take this problem into account, as considered by Rauch and Trindad (1999), Amurgo-Pacheco (2006), Rauch (1999), Soloaga and Winters (2001), Lederman and Özden (2004), Kirkpatrick and Watanabe (2005), Carrillo and Li (2002), Rose (2004a and 2005). However, Santos Silva and Tenreyro (2005) show that the multiplicative gravity specification

$$X_{ijt} = Y_{it}^{\beta_1} \cdot Y_{jt}^{\beta_2} \cdot N_{it}^{\beta_3} \cdot N_{jt}^{\beta_4} \cdot D_{ij}^{\beta_5} \cdot e^{\eta_{ij} + \sum_h \delta_h P_{ijth} + \sum_k \lambda_k F_{ijk} + u_{ijt}} \quad [6]$$

could be more appropriate. In more details, Santos Silva and Tenreyro (2005) consider different methods for comparisons (OLS, adding 1 to the trade variable, Tobit, NLS, PQML), but their tests systematically support the use of PQML technique (see section 3 note 33). Nevertheless, being the error term of the log-linearized specification heteroskedastic, then the statistical independence between the error term and the independent variables is violated and this leads to inconsistent estimates. Since the OLS and Tobit estimates are very similar, while the multiplicative model yields different results,[96] the authors ascribe these differences to the heteroskedasticity, rather than to truncation in data. The multiplicative specification has been estimated by using Nonlinear Least Square (Marquez-Ramos et al., 2006) or Poisson

[96] Results obtained by Santos Silva and Tenreyro (2005) show that the estimated parameters by the Poisson regression indicate that PTAs play a much smaller role. OLS estimates suggest that PTAs rise expected bilateral trade by 63 percent, whereas Poisson estimates indicate an average enhancement effect below 20 percent. OLS adding 1 to the trade variable and Tobit estimates generate extremely large and statistically significant coefficients for the trade-agreement dummy. The first method predicts that trade between countries that signed an agreement is on average 266 percent higher than that between countries without an agreement. The second predicts that trade between countries in PTAs is on average 100 percent larger. NLS predicts a significant and large effect for PTAs too, even if the other estimated parameters (i.e. GDP coefficients) are somewhat different.

Maximum Likelihood estimator (Westerlund and Wilhelmsson, 2006; Siliverstovs and Schumacher, 2006).

Another important econometric issue is the endogeneity of explanatory variables, for example GDP and the PTA variables. Under endogeneity of regressors a proper estimator to be employed is the instrumental variable (IV). IV method is considered by Eaton and Kortum (1997) to correct for endogeneity of the wage variable, by Tang (2005) for real exchange rate volatility, by Rose (2005) for the debt variable, by Rose (2000) for currency union and exchange rate volatility and by Tenreyro and Barro (2003) for common currency variable. To the best of my knowledge, only Agostino et al. (2007) carry out a test to verify the exogeneity of the PTA variable. Indeed, there could be a problem of simultaneity between trade flows and PTA variables, since it is not univocally determined if countries trade more because they are in a PTA or they belong to a PTA because they already traded relatively more with each other than with third countries. Another approach that could be used in order to take into account the endogeneity of PTA variables is the two-step IV method. In particular, according to Lederman and Özden (2004), the granting and removal of preference eligibility is a political decision. A country could be interested in trade with political allies or with strategically important countries (Lederman and Özden, 2004). This can be modelled using a treatment effect model, where the sample is divided into treated (the units that participate in a program, in our case the countries that belong to the PTA) and not treated (the countries that do not belong to the PTA), and the treatment (belonging to the PTA) is an endogenous process. Formally, the treatment issues is analogous to the sample selection problem. Consider the basic equation of primary interest [3]:

$$Y_1 = X_1'\beta_1 + \varepsilon_1 \qquad [7]$$

where now X_1 is only observed when d_2=1, that is when the unit participates in the treatment (that in our case consists in belonging to a PTA). Suppose that the model that explains the decision to join a PTA is the following:

$$d_2^* = Z_2'\gamma_2 + \varepsilon_2 \qquad [8]$$

with d_2=1 if the unit participates in the treatment (d_2^*>0) and d_2=0 otherwise ($d_2^* \le 0$).

Then, the conditional expectation of Y_i given d_2=1 is the following:

$$\begin{aligned} E\left(Y_{i,1} / d_2 = 1\right) &= E\left(Y_{i,1} / d_2^* > 0\right) = X_{i,1}'\beta_1 + E\left(\varepsilon_1 / d_2^* > 0\right) \\ &= X_{i,1}'\beta_1 + E\left(\varepsilon_1 / \varepsilon_2 > -Z_2'\gamma_2\right) \end{aligned} \qquad [9]$$

Similarly to the case of sample selection (eqs 3-5), if ε_1 and ε_2 are correlated, then estimates will be biased. Lederman and Özden (2004) and Baier and Bergstrand (2002 and 2005) address this issue by employing a two-step IV method, where in the first step a probit concerning the decision to join a PTA is modelled, and in the second step the gravity equation is estimated using the fitted probability obtained in the first step as an IV for the PTA variable.

Finally, a possibly more adequate specification would take into account the persistency of trade between country pairs. In fact, according to Bun and Klaassen (2002b, p.2), "for countries that traded a lot in the past, businesses have set up distribution and service networks in the partner country". In addition, consumers are familiar with the partner country's products (habit formation). Under these circumstances, therefore, it appears very likely that there would be relatively higher current bilateral trade between the countries. Hence, lagged trade could explain, at least in part, current trade, and ignoring this might well lead to incorrect inference. Dynamic specifications are estimated by using OLS (Freund and Weinhold, 2004; Gould, 1998; Mansfield and Bronson, 1997; Faruqee, 2004; Martinez-Zarzoso and Nowak-Lehmann, 2003); Least Square Dummy Variable Model (Bun and Klaassen, 2002a); Arellano-Bond (1991) estimator (Rose, 2004a; Micco et al., 2003); or GMM-system estimator (De Benedictis and Vicarelli, 2004; De Benedictis et al., 2005; Martinez-Zarzoso et al., 2006).

No paper among those reviewed considers jointly all these estimation issues. As far as PTA variables are concerned, only Emlinger et al. (2006) and Cipollina and Salvatici (2007) use a quantitative variable providing more information on the preferential schemes than dummies. But from an econometric point of view, they consider only the Heckman two-step procedure taking into account zero-trade flows. In this way, they disregard all the other potential kinds of bias.

To the best of my knowledge only Agostino et al. (2007) take care of relatively more empirical sources of bias. Indeed, they employ a fixed effects model which is robust to the presence of unobserved country heterogeneity. Moreover, they address the issue of non-random selection which zero-trade observations pose in the log-linearization of the gravity model and perform a test for the endogeneity of the preferential trade variable.[97] However, they disregard the trade persistency issue and do not consider an appropriate indicator of PTAs, which they measure by dummies and a multinomial variable (see section 3.5).

[97] Rose (2004a, 2004b and 2005) considers several different estimation methods to check for robustness of results. However, he does not consider jointly all sources of bias.

To sum up, we could be confident of the results obtained regarding the role of PTAs on trade only if all empirical issues are considered. A fully reliable empirical analysis should, firstly, be based on PTA variables adequately representing the policy instruments used and the preference granted. Secondly, a multiplicative gravity model with dynamic specification and fixed effects should be considered. Finally, instrumental variables should be used in order to correct for endogeneity of some regressors, and, more precisely, a two-step IV procedure to correct for endogeneity of the PTA variable should be employed.

5. Conclusions

The survey presented in this paper reviews contributions using gravity models to assess the effectiveness of PTAs in increasing trade and critically analyses the empirical choices made in these studies and their implications.

Papers differ in the aim pursued, in the gravity specification considered, in the sample used and in the estimation methods employed. Hence, comparisons among results must be made with extreme caution.

The main conclusion that emerges from this review is that PTAs tend to foster trade between the preferred country and that granting the preference. This conclusion is also reached in Cipollina and Salvatici (2006),[98] Greenaway and Milner (2002) and Nielsen (2003).[99]

In particular, the creation and the enlargement of the EU is found to enhance trade between both EU members and between EU and non-EU countries. A similar result is observed for EFTA. CMEA has been found to have a positive impact in Endoh (1999) and Gaulier et al. (2004), while Adam et al. (2003) obtain a negative impact. Furthermore, Adam et al. (2003) and Sissoko (2004) show that BAFTA is more effective than CEFTA. In addition, Breuss and Egger (1999) and Managi et al. (2005) find that NAFTA is more successful in increasing trade than EU; Nilsson (2005) obtains that EU trade policies in favour of LDCs have generated more trade than the trade agreements promoted by US. Results also show that PTAs among Asian, Latin and Central American, South Pacific and African

[98] In more details, the results of the meta-analysis show that the RTA effect is equivalent to an increase in trade by more than 11%. Furthermore, the impact tends to be larger in recent years, "and this could be a consequence of the evolution from "shallow" to "deep" trade agreements" (Cipollina and Salvatici, 2006, p. 20).

[99] Nielsen (2003) also shows that studies using partial equilibrium (PE) and general equilibrium (CGE) models generally obtain a positive and relatively small impact of PTAs on trade. As the author points out, "it is important to recognise that different modelling approaches should be seen as complements rather than substitutes" (Nielsen, 2003, p. 111). Indeed, according to Greenaway and Milner (2002, p. 12), "empirical gravity models can provide information for CGE models on bilateral trade elasticities with respect to incomes, prices and transaction cost barriers".

countries foster trade. As far as unilateral trade agreements granted by the EU are concerned, Persson and Wilhelmsson (2006) obtain a negative impact on trade of the PTA with ACP countries, while Nilsson (2002) and Oguledo and Macphee (1994) obtain a positive intra-trade effect. According to Nilsson (2002) the positive impact is not only due to the preferential regime but also to strong trading relations among EU and ACP countries dating back to colonial times. GSP granted by the EU has a positive effect on trade in Persson and Wilhelmsson (2006) and De Santis and Vicarelli (2006), and a negative one in Oguledo and Macphee (1994) and Verdeja (2006).

Negative intra-bloc impacts are obtained in numerous papers concerning different PTAs. This is a questionable result because there is no good reason to expect that a PTA should decrease trade between members (and, in fact, no good reason is provided to explain the result obtained).

These doubtful results could be due as well to empirical problems not adequately taken care of. One of these problems concerns the use of a dummy to proxy a PTA. This approach is problematic because a dummy captures all the factors specific to each country-pair and contemporaneous with the PTA. Moreover, it assumes that the impact of PTAs on trade is the same whatever the PTA.

Other empirical problems concern the estimation methods used. First of all, only some papers consider heterogeneity of countries. Since country specific effects are likely to be correlated with explanatory variables, then not considering this issue leads to biased estimates. Secondly, considering zero trade flows as missing values could produce biased estimates. In this sense, Santos Silva and Tenreyro (2005) show that a multiplicative gravity specification estimated by PQML is more adequate. Thirdly, endogeneity of explanatory variables and, in particular, of PTA variables must be considered. Finally, a proper specification should take into account also the persistency of trade between countries.

If adequate testing has not ruled out the presence of these estimation problems, then we cannot be confident of the results obtained regarding the effectiveness of PTAs on trade.

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