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The Trade Effects of Endogenous Preferential Trade Agreements†

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Structural new trade theory models have never been used to evaluate and quantify the role of preferential trade agreement (PTA) membership for trade in a way which is consistent with general equilibrium. Apart from filling this gap, the present paper aims at delivering an empirical model which takes into account both that PTA membership is endogenous and that the world matrix of bilateral trade flows contains numerous zero entries. These features are treated in an encompassing way by means of (possibly two-part) Poisson pseudo-maximum likelihood estimation with endogenous binary indicator variables in the empirical model. (*JEL* F11, F13, F15)

The unprecedented surge of preferential trade liberalization since World War II spurred theoretical and empirical work on the matter alike. Theoretical research illustrated under which conditions preferential trade agreements (PTAs) induce welfare gains for participants.[1] Econometric work confirmed that economic and political fundamentals determine preferential trade liberalization through PTA membership very much along the lines hypothesized by economic theory (see Scott L. Baier and Jeffrey H. Bergstrand 2002, 2004, 2009; Christopher S. Magee 2003; Hartmut Egger, Peter Egger, and David Greenaway 2008): PTAs are most likely concluded among large, similarly-sized, non-distant economies which have modern political systems. In part this empirical work has even strived for an identification of *causal effects* of PTA membership and found that, indeed, PTA membership causes bilateral trade.

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[1] The existing body of theoretical work on endogenous trade policy in general and endogenous PTA membership in specific is by far too large to be discussed here. However, we refer the interested reader to the excellent surveys by Dani Rodrik (1995), Richard E. Baldwin and Anthony J. Venables (1995), and Baldwin (2009), for details.

However, from a theoretical perspective, there are two major discomforts with seemingly all empirical work on the causal effects of PTA membership on trade flows. First, general equilibrium effects are ignored. All of the corresponding work relies on the so-called *stable unit treatment value assumption* (SUTVA) which requires that PTA membership only affects PTA insiders but outsiders not at all (see Jeffrey M. Wooldridge 2002; Colin A. Cameron and Pravin K. Trivedi 2005). Obviously, this is at odds with general equilibrium. James J. Heckman, Lance Lochner, and Christopher Taber (1998) emphasize and illustrate that treatment effects can be severely biased when ignoring general equilibrium effects. They criticize that the "*paradigm in the econometric literature on treatment effects is that…there are no spillovers…*" and argue that "*standard policy-evaluation practices are likely to be misleading…*" accordingly (p. 381). Second, the extensive margin of bilateral trade is forgotten about and sample selection is induced by focusing on log-transformed trade flows as outcome. This paper ventures for an alternative approach which pays explicit attention to both of these problems.

We pursue an empirical modeling strategy which is informed by three influential strands of recent empirical research in international economics: first, the work on empirical estimation of general equilibrium models where trade costs exert bilateral as well as multilateral effects on trade and GDP (see Jonathan Eaton and Samuel Kortum 2002; James A. Anderson and Eric van Wincoop 2003; Anderson 2009); second, research on zeros in bilateral trade matrices for any year or averages of years suggesting that the extensive margin of bilateral trade should be modeled explicitly in empirical analysis (see J. M. C. Santos Silva and Silvana Tenreyro 2006; 2008; and Elhanan Helpman, Marc Melitz, and Yona Rubinstein 2008); third, the literature on endogenous PTAs and their causal effects on trade flows (see Baier and Bergstrand 2002; 2007; 2009).[2] Interestingly, these obviously important three bodies of work are virtually unconnected.

This paper treats PTA membership as an endogenous determinant of bilateral trade while allowing for (numerous) zero bilateral trade flows in the empirical model, and respecting both the bilateral and multilateral effects of endogenous PTAs on trade in the quantification of PTA effects. In contrast to preceding work by Eaton and Akiko Tamura (1994); Santos Silva and Tenreyro (2006, 2008); and Helpman, Melitz and Rubinstein (2008), we allow (binary) determinants of exports to be endogenous. In particular, we suggest empirical models based on pseudo-maximum likelihood estimation with endogenous (binary) explanatory variables.

We apply these models to a cross-sectional dataset of bilateral trade flows and their determinants—among them a binary PTA membership indicator—for the year 2005. We compute cum-PTA bilateral trade flows and compare them to counterfactually predicted trade flows in a sine-PTA general equilibrium. Eliminating PTAs

[2] The quantification of the effects of preferential trade agreement (PTA) membership has been a major source of interest of empirical bilateral trade flow modelers for decades. See Jan Tinbergen (1962), Herbert Glejser (1968), Norman D. Aitken (1973), for some of the earliest examples and Caroline Freund (2000), Isidro Soloaga and L. Alan Winters (2001), and Céline Carrère (2006) for more recent ones. Greenaway and Chris Milner (2002) provide a useful survey. For decades, the dominant paradigm in related work was that countries were randomly assigned to PTAs. Only recently, Baier and Bergstrand (2002, 2004, 2007, 2009), Magee (2003), and Egger, Egger, and Greenaway (2008) allowed for PTAs to be endogenous to trade in an econometric sense.

reduces trade flows among members directly, but it also entails indirect effects on third countries through the impact of PTAs on producer prices, consumer prices, and GDP.

Our findings may be summarized as follows. The results shed light on three potentially large biases associated with the ignorance of the three mentioned issues: general equilibrium (third-country) effects of PTA membership; zeros in trade matrices; and the endogeneity of PTAs. The biases are of different magnitude, though. Let us use the PTA-related change of bilateral exports among members relative to nonmembers to quantify biases. Then, for instance, a log-linear model of exports which ignores general equilibrium effects on top of the other problems leads to a bias of -186 percentage points or -79 percent relative to the preferable two-part approach (which inter alia controls for the presence of heterogeneous firms). A one-part Poisson pseudo-maximum likelihood model which disregards non-random selection into positive exports and treats PTA membership as exogenous leads to a bias of the impact of PTAs on members' relative to nonmembers' trade by -176 percentage points or -75 percent relative to a two-part model which copes with all of the mentioned problems. A one-part model which acknowledges endogenous PTA membership but disregards the problem of an excessive number of zeros in the data leads to a downward bias of the PTA effect by about -73 percentage points or -31 percent.

Apart from these biases there are another two which seem relatively less important. Disregarding the presence of heterogeneous firms appears less relevant than the mentioned biases. A two-part model without heterogeneous firm controls leads to a bias of the average estimated PTA effect of -15 percentage points or -6 percent in general equilibrium. In comparison, it is even less harmful to ignore that PTA membership effects are heterogeneous due to the variation in most-favored nation tariff rates. Ignoring heterogeneous tariffs in the preferable two-part PTA model leads to a downward bias of the PTA-induced effect of less than one-tenth of a percentage point.

The remainder of the paper is organized as follows. The next section briefly introduces the bilateral trade flow model we will rely upon. Section II points out three problems with the implementation of that model in applied work targeted towards the analysis of PTA membership effects on trade. Section III describes the specification and data. Section IV introduces the modeling strategy to overcome these obstacles by treating zero trade flows implicitly, and presents the corresponding estimation results. Section V derives a zero-inflated gravity equation, lays out econometric two-part models, and summarizes the estimation results thereof. Section VI computes the impact of PTA membership as observed in the year 2005 and compares it to an unobserved counterfactual situation without any PTA memberships in the same year. The last section concludes with a summary of the most important findings.

I. Specifying Bilateral Trade Flows in the Vein of Anderson and van Wincoop (2003)

Anderson and van Wincoop (2003) derive a general representation of bilateral aggregate nominal trade flows in new trade theory models with one sector and N countries. For instance, such models include the ones of Anderson (1979) or Paul R. Krugman (1980) with love-of-variety preferences à la Avinash K. Dixit and Joseph

E. Stiglitz (1977). Their framework can be briefly introduced as follows. Let us denote nominal exports of country i to country j (with $i, j = 1, \ldots, N$) by X_{ij} and refer to trade costs associated with exports from country i to j as t_{ij}. Finally, use y_i, y_j, and y_W for country i's, country j's, and world GDP (total expenditures), respectively. Then, nominal bilateral exports are determined as

$$X_{ij} = \frac{y_i y_j}{y_W} t_{ij}^{1-\sigma} \Pi_i^{\sigma-1} P_j^{\sigma-1}, \tag{1}$$

where σ is the elasticity of substitution among products (variants) and Π_i, P_j are so-called *multilateral resistance* (MR) terms for exporters and importers, respectively. MR terms reflect multilateral (non-linearly weighted) trade costs firms of an exporting country and consumers in an importing country are faced with. Empirically, these MR terms are not observed but they can be readily derived as implicit solutions of the following set of $2N$ equations[3]

$$\Pi_i^{1-\sigma} = \sum_{j=1}^{N} \left(t_{ij}^{1-\sigma} P_j^{\sigma-1} y_j/y_W\right); \quad P_j^{1-\sigma} = \sum_{i=1}^{N} \left(t_{ij}^{1-\sigma} \Pi_i^{\sigma-1} y_i/y_W\right) \forall i, j. \tag{2}$$

The structural representation of the model brings about a substantial advantage over other, reduced-form (and partly ad-hoc) specifications of gravity models of bilateral trade. Heckman, Lochner, and Taber (1998, p. 381) mention that "*standard policy-evaluation practices are likely to be misleading*" if individual (in our case, country-pair specific) choices affect others' economic outcome, as is the case in general equilibrium models like the one we are considering. "*The paradigm in the econometric literature on treatment effects is that* (…) *there are no spillovers* (…)." Since spillover effects from one country-pair to others are at the very heart of the matter, a full account of the impact of trade costs or PTA membership on exports in general equilibrium needs to respect their effect on all variables on the right-hand side of (1): on trade costs as such (t_{ij}), on exporter GDP (y_i), importer GDP (y_j), and world GDP (y_W), respectively (since they are a function of trade flows), and on the exporter and importer MR terms (Π_i and P_j), respectively. Notice that the direct effects of trade costs are generally dampened by the MR terms as illustrated in Anderson and van Wincoop (2003).

Since direct measures of trade frictions t_{ij} are typically not available, one uses proxy variables thereof. The bilateral log-distance between countries' capitals ($DIST_{ij}$), a common international border indicator ($BORD_{ij}$), and a common official language indicator ($LANG_{ij}$) are typical examples. In most empirical models of bilateral trade flows, trade policy is accounted for as an element of t_{ij} by including an indicator variable of preferential trade agreement membership (PTA_{ij}). The commonly adopted assumption about the relationship between t_{ij} and these proxy variables is

$$t_{ij}^{1-\sigma} = \exp(\beta_1 DIST_{ij} + \beta_2 BORD_{ij} + \beta_3 LANG_{ij} + \cdots + \delta PTA_{ij}). \tag{3}$$

[3] Notice that the $2N$ equations have to be properly normalized to avoid multiple solutions to the system of $2N$ equations (see Anderson 2009).

Substituting (3) into (1), we obtain the multiplicative model

$$X_{ij} = \exp(\mathbf{Z}_{ij}'\boldsymbol{\beta} + \delta PTA_{ij} + \alpha_i + \gamma_j), \tag{4}$$

where $\mathbf{Z}_{ij} = (1, DIST_{ij}, BORD_{ij}, \ldots)'$ is a vector containing a constant and all trade cost or trade facilitating variables except PTA_{ij}. Moreover, $\beta = (\beta_0, \beta_1, \beta_2, \ldots)'$ is a vector of coefficients corresponding to the elements in $\mathbf{Z}_{ij}$. $\alpha_i = \ln(y_i \Pi_i^{\sigma-1})$ and $\gamma_j = \ln(y_j P_j^{\sigma-1})$. In this model, the coefficient on the constant is defined as $\beta_0 = -\ln y_W$.

II. Empirical Problems with the Implementation of a Structural Gravity Model

Anderson and van Wincoop (2003) suggest estimating a stochastic version of (4)

$$X_{ij} = \exp(\mathbf{Z}_{ij}'\boldsymbol{\beta} + \delta PTA_{ij} + \alpha_i + \gamma_j)\epsilon_{ij}, \tag{5}$$

by taking the logs of both the left-hand-side and the right-hand-side and essentially minimizing the sum of squared residuals subject to (2). Alternatively, the parameters β and δ can be estimated directly by treating α_i and γ_j as fixed country effects. Given these parameters, the $2N$ multilateral resistance terms in (2) may be computed subsequently. Since general equilibrium effects are fully captured by the country fixed effects, estimation of β and δ does not hinge upon the general equilibrium structure of the model. In fact, it is well-known that the econometric specification (5) can represent a wide range of (one-sector) models including the multi-country version of the Dixit-Stiglitz-Krugman model, Eaton and Kortum (2002), or Robert C. Feenstra (2004). Hence, most of what we will talk about with regard to estimation below applies to a wide range of empirical models that are informed by general equilibrium theory. The choice of the underlying theoretical model determines the magnitude and transmission channels of comparative static effects but not parameter estimates.

With parameter estimation, two issues may arise in such an empirical context. First and most importantly, recent work in international trade emphasizes that PTA membership should be treated as an *endogenous* rather than an *exogenous* determinant of trade (see Baier and Bergstrand 2002, 2007, 2009; Magee 2003). Baier and Bergstrand (2004) derived theoretical hypotheses about the determinants of PTA membership which work well in empirical applications. Yet, while previous work put great effort into identifying the causal effects of (endogenous) PTA membership, the empirical paradigm has been using microeconometric methods for program evaluation which prevent structural estimation of the impact of PTA membership as suggested by equations (1) and (2).[4] This research thus assumed that PTA membership of one country-pair only affects this pair's bilateral exports but not those of other country-pairs. The latter feature is at odds with both intuition and general equilibrium. We will show how the model in 1 can be adapted to account for some

[4] Previous work predominantly relied on Heckman-type switching regression models (Baier and Bergstrand 2002; Magee 2003) or matching methods based on the propensity score (Baier and Bergstrand 2002, 2009).

endogenous trade frictions, still obeying (2). Obviously, such a goal can only be achieved by means of instrumental variable estimation.

Second, depending on the dataset in use, the $N(N-1)$-size vector X of bilateral exports with typical element X_{ij} may contain numerous zeros (see Helpman, Melitz, and Rubinstein 2008) whose omission (by taking the log of the left-hand-side of the model) would in general lead to an efficiency loss and to inconsistent parameter estimates. Some authors have circumvented the problem of omitting zero trade flows by adding a small positive constant to X, a transformation that enables logarithmizing all X_{ij}. Santos Silva and Tenreyro (2006) show that this approach leads to inconsistent parameter estimates as well. The severity of the bias resulting from this ad-hoc solution can be quite large. Thus, estimating the model in its original multiplicative form (5) seems highly preferable. Furthermore, multiplicative models as in (5) imply by construction that higher conditional expectations go hand in hand with higher conditional variances. This pattern of heteroskedasticity is a well-known stylized fact of trade data, rendering multiplicative estimation of the model even more attractive.

We elaborate on these issues in Section IV, where we present an econometric model of the gravity equation which is able to appropriately deal with both of these problems. Before that, we describe our general specification and the data used.

III. Specification and Data

We broadly follow Baier and Bergstrand (2004) and Egger, Egger, and Greenaway (2008) to model selection into PTA membership as a function of three sets of characteristics: variables capturing political affinities or impediments to bilateral trade liberalization, proxies for iceberg trade costs, and country size and relative factor endowments. We classify two countries as belonging to a common PTA, if they were active since 2005 or earlier as notified to the World Trade Organization. The data were augmented and corrected by using information from PTA secretariat webpages and they were compiled to obtain a binary dummy variable reflecting PTA memberships for the year 2005. The three sets of exogenous variables contain the following elements:

Variables Capturing Political Affinities or Impediments to Bilateral Trade Liberalization.—Political scientists have pointed to a number of political factors which are hypothesized to affect bilateral trade flows (see Egger, Egger, and Greenaway 2008, for a brief survey). The corresponding variables reflect characteristics of political systems. The associated variables are based on the data collected in the Polity IV Project (see Monty G. Marshall and Keith Jaggers 2007). In particular, we include the absolute difference in a score variable, measuring the autocracy of an exporter and an importer, respectively ($AUTOC_{ij}$);[5] the absolute difference in a variable, measuring the durability of an exporter's and an importer's political regime,

[5] $AUTOC_{ij}$ measures Institutionalized Autocracy in a country. In the most extreme form, autocracy suppresses competitive political participation, chief executives are chosen within a small political elite, and once in office exercise power almost without institutional constraints. The source data vary between 0 and 98.

respectively ($DURAB_{ij}$);[6] the absolute difference in a score variable, measuring the political competition in the government of an exporter and an importer, respectively ($POLCOMP_{ij}$).[7]

Proxies for Iceberg Trade Costs.—Log bilateral (great circle) distance between two countries' capitals ($DIST_{ij}$);[8] an indicator variable which is one in case of a common land border between countries i and j and zero else ($BORD_{ij}$); an indicator variable which is set to one if two countries have a common language and zero else ($LANG_{ij}$); an indicator variable which is set to one if two countries are located at the same continent and zero else ($CONT_{ij}$); an indicator variable which is set to one if one of two countries had been a colony of the other in the past and zero else ($COLONY_{ij}$); an indicator variable which is set to one if one of two countries had been a colony of the other after the year 1945 and zero else ($CURCOL_{ij}$); an indicator variable which is set to one if the two countries had a common colonizer in the past and zero else ($COMCOL_{ij}$); an indicator variable which is set to one if one country was part of the other in the past and zero else ($SMCTRY_{ij}$). All of the mentioned trade cost indicators are taken from the geographical database provided by the Centre d'Études Prospectives et d'Informations Internationales (CEPII). The list of variables in Baier and Bergstrand (2004) did not include (country dummies and) $LANG_{ij}$, $COLONY_{ij}$, $CURCOL_{ij}$, $COMCOL_{ij}$, or $SMCTRY_{ij}$.

Country Size and Relative Factor Endowments.—Exporter and importer country size in terms of their log GDP as two separate determinants as well as all other country-specific determinants such as population, capital-labor ratio, etc., are fully accounted for by fixed exporter and importer dummy variables. Baier and Bergstrand (2004) use non-linear transformations of exporter and importer log GDP and include log total bilateral GDP and log similarity of bilateral GDP as determinants of PTA. In addition, they include two measures of relative factor endowment differences. One of them reflects the capital-labor relative factor endowment difference between two countries in a pair and the other one captures the capital-labor relative factor endowment difference between that pair and the rest of the world. We recreated these four variables, although for reasons of data availability (the dataset used here contains 15,750 country-pairs while the one in Baier and Bergstrand 2004, covered only 1,453 country-pairs) we had to use real GDP per capita instead of employing capital-labor ratios.[9] However, capital-labor ratios are highly correlated with real GDP per capita. While these variables worked well for the determination of PTA membership, we ran into convergence problems when using these jointly for the PTA and exports equations. We conducted robustness checks using two of them at a time (not reported) to make sure that our reported results where not changed by their

[6] $DURAB_{ij}$ measures the number of years since the most recent regime change or the end of a transition period without any stable political institutions in place. $DURAB_{ij}$ is computed for all years beginning with the first regime change since 1800 or the date of independence if that event occurred after 1800.

[7] $POLCOMP_{ij}$ measures to which degree party participation is regulated in a country and to which degree there is competition in participation. The source data vary between 0 and 98.

[8] Baier and Bergstrand (2004) include a variable which is defined as $NATURAL_{ij} = -DIST_{ij}$. Hence the expected sign of $DIST_{ij}$ is exactly the opposite of the one of $NATURAL_{ij}$.

[9] Data on real GDP and population are taken from the World Bank's World Development Indicators.

Table 1—Descriptive Statistics of the Data

Variable (1)	Description (2)	Mean (3)	SD (4)	Min. (5)	Max. (6)
X_{ij}	Nominal exports in million US dollars	305.9274	3,257.2670	0	213,763.06
$\mathcal{I}_{ij}$	Indicator variable taking value one if $X_{ij} > 0$	0.6280	0.4834	0	1
PTA_{ij}	Indicator variable taking value one if two countries belong to a common PTA since 2005 or earlier	0.2226	0.4160	0	1
$DIST_{ij}$	Log distance	8.2002	0.8267	3.2467	9.4191
$BORD_{ij}$	Common border indicator variable	0.0210	0.1432	0	1
$LANG_{ij}$	Common language/ethnicity indicator variable	0.1393	0.3463	0	1
$COLONY_{ij}$	Colony indicator variable	0.0152	0.1225	0	1
$COMCOL_{ij}$	Common colonizer indicator variable	0.0777	0.2677	0	1
$CURCOL_{ij}$	Colony after 1945 indicator variable	0.0084	0.0912	0	1
$SMCTRY_{ij}$	Same country indicator variable	0.0088	0.0935	0	1
$CONT_{ij}$	Same continent indicator variable	0.2303	0.4211	0	1
$RGDPsum_{ij}$	Log of sum of real GDPs	25.2322	1.8080	19.9296	30.1824
$RGDPsim_{ij}$	Similarity of real GDPs	−2.1131	1.4877	−9.7690	−0.6931
DKL_{ij}	Difference between log of capital-labor relative factor endowments between pair *ij*	1.8217	1.2944	0.0001	6.1001
$DROWKL_{ij}$	Difference between log of capital-labor relative factor endowment between pair *ij* and rest of the world	1.4852	0.6493	0.0659	3.7327
$DURAB_{ij}$	Durability of an exporter's and an importer's political regime	29.4047	29.2178	0	100
$POLCOMP_{ij}$	Political competition index	8.8961	19.9440	0	98
$AUTOC_{ij}$	Autocracy index	7.9867	18.9474	0	98
Observations			15,750		

inclusion or omission. In the specification used in the following sections, we omitted these variables, as the detected patterns for parameter estimates were not affected by their inclusion or omission in any substantial way.

Finally, our empirical model includes the following trade cost variables in $\mathbf{Z}_{ij}$ in the nominal exports outcome equation (5): $DIST_{ij}$, $BORD_{ij}$, and $LANG_{ij}$, $CONT_{ij}$, $DURAB_{ij}$, $AUTOC_{ij}$, $POLCOMP_{ij}$, $CURCOL_{ij}$. Otherwise, nominal exports are a function of a complete set of exporter and importer dummy variables,[10] and of (potentially endogenous) PTA_{ij}. Data on bilateral exports in nominal US dollars are collected from the United Nation's World Trade Database.

Table 1 summarizes mean, standard deviation, minimum and maximum of the distribution of the dependent and independent variables employed in the estimated models. Here, we would like to emphasize that about 37 percent of the cells of the

[10] Which capture GDP and MR terms in (5).

bilateral exports matrix are zero and about 22 percent of the 15,750 country-pairs in our dataset are members of a common PTA.

IV. Estimating a Gravity Model with Zero Export Flows and Endogenous PTA Membership

For an assessment of the effects of PTA membership on trade flows, it is necessary to obtain consistent estimates of the unknown parameter vector β and the PTA parameter of interest, δ. However, δ does only reflect direct effects of PTA membership on exports. To quantify total effects—which also account for feedback across countries consistent with general equilibrium—we need to compute counterfactual exports without PTA membership. The latter also account for the impact of PTA membership on GDPs and MR terms as explained in Section I. We will quantify the impact of PTA memberships by comparing predicted exports of PTA insiders with PTAs as of 2005 relative to outsiders with predicted relative trade flows in a counterfactual scenario without any PTAs. While the corresponding simulation results are presented in Section VI, our first objective in the subsequent sections is to consistently estimate β and δ.

A. *Econometric Model*

Since the parameters of interest in model (5) are β and δ, terms α_i and γ_j can be considered, from an econometric point of view, as nuisance parameters. The model to be estimated thus represents a two-way country-specific effects model, where α_i and γ_j subsume the effects of GDP and MR terms, but may depend on other country-specific factors as well. The appropriate econometric methods depend on the assumptions on the relationship between (α_i, γ_j) and the regressors, $\mathbf{Z}_{ij}$ and PTA_{ij}. If (α_i, γ_j) were independent of $\mathbf{Z}_{ij}$ and PTA_{ij}, random effects estimation would be consistent and efficient. However, the underlying economic model suggests that α_i and γ_j depend on $\mathbf{Z}_{ij}$ and PTA_{ij}. Therefore, the model should be treated as a two-way fixed effects model.

There are two important differences to a standard panel data model, though. First, this model is non-linear, and simple (within) transformations to eliminate the fixed effects are not available. Second, since the data consist of all possible pairs of N countries, and each country is observed as both exporter and importer, there are $N(N-1)$ observations. Hence, adding one country to an existing set of N economies gives $2N$ additional observations but only 2 additional parameters. It follows that there is no incidental parameter problem, and the country fixed effects can be estimated consistently (for $N \to \infty$) by including a dummy variable for each importer and exporter country.[11] This procedure is computationally intensive, given

[11] The classical incidental parameter problem in non-linear panel models says the following. Suppose that data vary in two dimensions, one of which is small (with a fixed number of T units) and one is large (with $N \to \infty$ units). Then, it is impossible to estimate individual fixed effects for each unit in N consistently. Similarly, the slope parameters of covariates can then not be estimated consistently.

the large number of $2N - 2$ fixed effects to be estimated, but it is straightforward in its application.

The conditional expectation function (CEF) of model (5) is

$$E(X_{ij} \mid \mathbf{Z}_{ij}, PTA_{ij}, \alpha_i, \gamma_j) = \exp(\mathbf{Z}'_{ij}\beta + \delta PTA_{ij} + \alpha_i + \gamma_j)\mathrm{E}(\epsilon_{ij} \mid \mathbf{Z}_{ij}, PTA_{ij}, \alpha_i, \gamma_j). \tag{6}$$

Under the assumption of exogenous PTA membership, $\mathrm{E}(\epsilon_{ij} \mid \mathbf{Z}_{ij}, PTA_{ij}, \alpha_i, \gamma_j) = 1$ and model (5) would be simply an exponential CEF model. However, acknowledging that PTA membership is potentially endogenous, we want to allow for possible correlation between the error term ϵ_{ij} and the propensity to form an agreement. To tackle this problem we implement an instrumental variable method based on the joint distribution of ϵ_{ij} and PTA_{ij}. Specifically, assume the following reduced-form equation for PTA_{ij},

$$PTA_{ij} = \begin{cases} 1 & \text{if} \quad \mathbf{W}'_{ij}\theta \geq v_{ij}, \\ 0 & \text{if} \quad \mathbf{W}'_{ij}\theta < v_{ij}, \end{cases} \tag{7}$$

where $\mathbf{W}_{ij}$ is a vector comprised of variables affecting a country i's participation decision in a preferential trade agreement with country j. The elements of $\mathbf{W}_{ij}$ have been listed in Section III and they contain all elements of $\mathbf{Z}_{ij}$ as well as instrumental variables excluded from (6). Endogeneity arises if the errors v_{ij} and ϵ_{ij} are not statistically independent. Following Joseph V. Terza (1998), it is possible to derive a tractable form of $E(X_{ij} \mid \mathbf{Z}_{ij}, PTA_{ij}, \mathbf{W}_{ij}, \alpha_i, \gamma_j)$ under the assumption of bivariate normality of v_{ij} and $\ln(\epsilon_{ij})$, which leads to the following expressions

$$E(X_{ij} \mid \mathbf{Z}_{ij}, PTA_{ij}, \mathbf{W}_{ij}, \alpha_i, \gamma_j) = \lambda_{ij} \, \Psi_{ij}, \tag{8}$$

with

$$\lambda_{ij} \equiv \exp(\mathbf{Z}'_{ij}\beta + \delta PTA_{ij} + \alpha_i + \gamma_j) \quad \text{and} \tag{9}$$

$$\Psi_{ij} \equiv E(\epsilon_{ij} \mid \mathbf{Z}_{ij}, PTA_{ij}, \mathbf{W}_{ij}, \alpha_i, \gamma_j)$$

$$= PTA_{ij} \frac{\Phi(\vartheta + \mathbf{W}'_{ij}\theta)}{\Phi(\mathbf{W}'_{ij}\theta)} + (1 - PTA_{ij}) \frac{1 - \Phi(\vartheta + \mathbf{W}'_{ij}\theta)}{1 - \Phi(\mathbf{W}'_{ij}\theta)}.$$

The last equality follows from joint normality of the errors, where $\Phi(\cdot)$ denotes the cumulative distribution function of the standard normal distribution.[12] The

[12] Note that the assumption of normality leads to a Probit model for PTA_{ij} as is common in the empirical literature. As for $\ln(\epsilon_{ij})$, which is an additive element to the linear index $\mathbf{Z}'_{ij}\beta + \delta PTA_{ij} + \alpha_i + \gamma_j$, it can be thought of as unobserved heterogeneity stemming from omitted variables. Assuming normality here does not seem wholly unreasonable, since a case can be made for normality even if some omitted variables are not normally distributed, as their sum would tend to be so by some version of the central limit theorem if only the omitted variables were sufficiently numerous and independent.

parameter ϑ is equal to the square root of the variance of $\ln(\epsilon_{ij})$, multiplied by ρ, the correlation coefficient between v_{ij} and $\ln(\epsilon_{ij})$. If $\rho = 0$, the errors are independent, and $\Psi_{ij} = 1$ so that the conditional expectation of X_{ij} in (8) simplifies to λ_{ij}, which is exactly the special case considered in (6) with $E(\epsilon_{ij} | \mathbf{Z}_{ij}, PTA_{ij}, \alpha_i, \gamma_j) = 1$. However, if $\rho \neq 0$, estimation of the parameters β contained in λ_{ij} will be inconsistent if Ψ_{ij} is neglected.[13]

The recent literature has suggested non-linear least squares (NLS) as well as various pseudo-maximum likelihood (PML) estimators as the preferred approaches to estimate multiplicative gravity models such as (6) with $E(\epsilon_{ij} | \mathbf{Z}_{ij}, PTA_{ij}, \alpha_i, \gamma_j) = 1$ (Santos Silva and Tenreyro 2006).[14] These estimators differ in their weighting functions, and thus in efficiency. Santos Silva and Tenreyro (2006, 2008) show that if the conditional variance of the exports is proportional to the conditional mean, then the first-order conditions from minimizing the squared errors of the model are numerically equivalent to the first-order conditions of the Poisson PML model. Also, they find that the Poisson PML estimator performs well compared to other PML and NLS estimators in a series of different Monte Carlo simulation setups.[15]

Likewise, the parameters of model (8) can be estimated by non-linear least squares, by minimizing the sum of squares of $(X_{ij} - \lambda_{ij}\Psi_{ij})$ as in Terza (1998), or by Poisson PML estimation where the conditional expectation is now $\lambda_{ij}\Psi_{ij}$. As before, the NLS estimator gives more weight to observations with larger trade flows, while the Poisson PML estimator gives equal weight to all observations. While both techniques yield consistent estimates of the parameters if the conditional mean (8) is correctly specified, the results reported in Santos Silva and Tenreyro (2006) strongly encourage us towards viewing Poisson PML estimates as more efficient. The Poisson PML estimator for the model with endogenous PTA is implemented as a two-step estimator. In a first step, consistent estimates of θ are obtained from a Probit regression of model (7). In a second step, we replace θ by $\hat{\theta}$ in (8) and use Poisson PML to estimate the remaining parameters β, δ and ϑ. Second-step standard errors have been adjusted to account for the variance of first-step estimates.

B. *Estimation Results*

Using the data described in Section III, we estimated the parameters of the structural models of Section IVA by Poisson PML. Table 2 displays estimates of three alternative models of nominal bilateral exports in US dollars (X_{ij}). Column 2 summarizes estimates from a naïve log-linear model which simply drops all data points for which bilateral exports are zero and treats PTA membership as exogenous.

[13] An alternative estimation technique which does not rely on bivariate normality is the GMM approach of F. A. G. Windmeijer and Santos Silva (1997). However, no comparable extension of this GMM approach for the two-part model has been proposed. Our parametric assumptions allow us to extend the estimator to the two-part model of Section VB.

[14] Under PML the information matrix equality does not hold and robust standard errors are computed using the "sandwich" estimator involving the inverse of the Hessian matrix and the outer product of the gradient.

[15] Alternative consistent estimators include other members of the linearized exponential family such as the normal PML model or the gamma PML model. The negative binomial model is not a linear exponential family unless the dispersion parameter is set to an arbitrary constant. We do not report such other estimates since, with our data and specifications, their numerical properties were poor and convergence was hard to achieve, the likely reason being the presence of numerous fixed effects to which the Poisson PML estimator appears to be less sensitive to.

Table 2—Estimation Results for Gravity Models for Trade

	Exogenous PTA		Endogenous PTA	
Regression	$E(\ln(X_{ij}))$	$E(X_{ij} \mid \cdot)$	$\Pr(\mathrm{PTA}_{ij} = 1 \mid \cdot)$	$E(X_{ij} \mid \cdot)$
Estimator	OLS	Poisson PML	Probit ML	Poisson PML
(1)	(2)	(3)	(4)	(5)
PTA	0.3943	0.5548	—	1.1471
	(0.0499)	(0.1256)		(0.3838)
DIST	−1.2342	−0.4998	−1.0737	−0.3971
	(0.0417)	(0.0492)	(0.0403)	(0.0675)
BORD	0.7264	0.7263	−0.4687	0.7405
	(0.1115)	(0.0726)	(0.1292)	(0.0748)
LANG	0.6490	0.1553	−0.1193	0.2079
	(0.0555)	(0.0813)	(0.0626)	(0.0698)
CONT	−0.0143	0.2736	0.7650	0.1506
	(0.0549)	(0.1222)	(0.0479)	(0.1579)
DURAB	−0.0026	−0.0038	−0.0072	−0.0041
	(0.0006)	(0.0009)	(0.0010)	(0.0009)
POLCOMP	0.0106	0.0737	−0.0483	0.1044
	(0.0010)	(0.0327)	(0.0091)	(0.0271)
AUTOC	−0.3784	−0.1039	0.0480	−0.1391
	(0.0122)	(0.0325)	(0.0098)	(0.0291)
CURCOL	1.3321	0.7246	0.5189	0.6179
	(0.1402)	(0.1695)	(0.2468)	(0.1833)
COLONY	—	—	0.1356	—
			(0.1941)	
COMCOL	—	—	0.5519	—
			(0.0719)	
SMCTRY	—	—	1.2275	—
			(0.2496)	
$\hat{\vartheta}$	—	—	—	−0.3708
				(0.1810)
F-stat.	—	—	88.7790	2.4949
p-value of *F*	—	—	0.0000	0.4762
F-stat. OIR	—	—	—	1.4530
p-value of OIR	—	—	—	0.4836
Observations	9,891	15,750	15,750	15,750
Countries	126	126	126	126

Notes: The sources of the data are the United Nations' World Trade Database and the Centre d'Etudes Prospectives et d'Informations Internationales. All regressions include importer and exporter fixed effects. The last two columns indicate results where PTA was instrumented. $\hat{\vartheta}$ is a measure for potential endogeneity of PTA_{ij}. "*F*-stat." and "*p*-value of *F*" refer to a test of joint significance of COLONY, COMCOL and SMCTRY in the respective equation. The *F*-statistic should be significantly different from zero in column 4 but not in column 5. "*F*-stat. OIR" and "*p*-value of OIR" refer to a test for over-identifying restrictions in the corresponding log-linear IV model.

Obviously, the corresponding estimates in column 2 differ quite starkly from the ones in columns 3 and 5 and so do their comparative static effects so that we will dismiss this estimator and not refer to it further in the subsequent discussion. Column 3 reports parameters and robust standard errors of a Poisson PML model that treats PTA_{ij} as exogenous. In column 5, PTA_{ij} is allowed to be endogenous. For this, we used a first-stage probit model based on the covariates mentioned in Section III and summarized in column 4 of Table 2. In principle, the model for endogenous PTAs does not need instruments to be identified if the distributional assumptions are met. As we did not wish to rely on functional form alone, we excluded a subset of the first-stage variables $\mathbf{W}_{ij}$ from the set of second stage variables $\mathbf{Z}_{ij}$ to act as instruments.

The instruments should have an effect on the probability to form a PTA, but they should not have other, direct effects on exports. The first assumption can be tested by performing an F-test on the joint relevance of the instruments in the reduced-form equation in column 4. We test the second assumption in two ways. On the one hand, we include them as additional regressors in the outcome equation and test for their relevance on outcome beyond their role for PTA. We do so by performing F-tests on the joint relevance of the instruments in the model in column 5. Moreover, we test whether the instruments pass a conventional test for overidentifying restrictions in a log-linearized version of the model for positive exports. The variables *COLONY*, *COMCOL*, and *SMCTRY* are significant determinants of PTA as can be seen from the low p-value of the reported F statistic in column 4, and they pass the tests for instrument validity according to the insignificant F-tests on overidentifying restrictions in column 5. Hence, countries are more likely to select into PTA membership given a shared colonial past, but—after controlling for other determinants of trade flows—these determinants do not directly affect trade.[16]

The results in Table 2 suggest the following conclusions. First of all, selection into PTAs based on observables is positive for some variables such as *CONT* and *CURCOL*: these factors raise the probability of joining a PTA and also tend to have a trade-increasing effect, implying that particularly those country-pairs which display a high level of goods trade flows select into PTAs anyway. Notice that this result is consistent with the hypothesis in Baier and Bergstrand (2004) according to which PTAs exhibit the highest welfare gains in countries where bilateral trade flows would be (and are) large. However, there are also variables which negatively affect both selection into PTAs as well as bilateral exports: examples are *DIST* and *DURAB*. Finally, there are observables which have the opposite impact on selection into PTAs and on exports: for instance, *BORD*, *LANG*, and *POLCOMP*, affect selection into PTAs negatively but exports positively; on the contrary, *AUTOC* affects selection into PTAs positively but exports negatively.

Second, there is evidence for selection into PTAs on unobservables. Endogeneity of PTA_{ij} can be assessed by a simple t-test on $\hat{\vartheta}$, an estimate of the (scaled) correlation between PTA_{ij} and the stochastic error in the exports. If PTA_{ij} is exogenous, the correlation must be zero, so that the null hypothesis $\vartheta = 0$ provides a valid test for exogeneity. We find that $\hat{\vartheta}$ is negative and significant, thus rejecting exogeneity of PTA_{ij}. A negative ϑ indicates that unobservables (i.e., bilateral factors other than the economic and politic determinants which we include in our models) favoring the creation of a PTA on average come along with unobservables that have a negative impact on bilateral trade. This negative self-selection based on unobservables leads to a downward bias in the estimated PTA parameter: the point estimate increases as we abandon the assumption of exogeneity. The remaining parameters are fairly stable across columns 3 and 5.

[16] In contrast to this, more recent colonial ties (*CURCOL*, which measures colonizing after 1945) do have a significant impact on trade beyond influencing PTA membership.

The results from the probit estimation for the reduced-form equation of PTA are broadly in line with comparable previous work. The political variables *DURAB*, *POLCOMP*, and *AUTOC* turn out to be important for the decision to join a PTA as in Egger, Egger, and Greenaway (2008).[17] These variables were not included in the models of Magee (2003) or Baier and Bergstrand (2004), nor were the colonial history variables. However, the finding of statistically significant effects of a positive influence if countries are on the same continent (*CONT*) is consistent with Baier and Bergstrand (2004), and the negative effect of distance is in line with both Baier and Bergstrand's (2004) cross-sectional analysis and Magee's (2003) cross-section and panel results.

V. Modeling Zero Trade Flows Explicitly

In our dataset, 37.2 percent of all entries in the bilateral trade matrix were zeroes. The previous approach accommodated those zero trade flows implicitly. We did not need to exclude non-trading country-pairs, nor did we artificially change the source data (e.g., by adding a positive constant to all export flows as in Gabriel J.Felbermayr and Wilhelm Kohler 2006) to allow for log-linearization. Yet the high incidence of zero trade makes it potentially interesting to split the overall effect of PTA membership on expected trade volumes into its two component parts, an effect at the extensive country margin of exports—i.e., the number of pairings which started exporting because of PTA membership—and the effect at the the intensive country margin—the extent to which PTA membership raised exports among pairs that traded already.

The focus on country margins is different from the recent emphasis on firm margins introduced by Helpman, Melitz, and Rubinstein (2008). In their model, trade volumes (our intensive country margin) can change due to new (and less productive) firms entering into export markets, or due to existing firms expanding their activities. For our analysis, this distinction is unimportant, as both effects are part of the causal pathway from PTA membership to increased country trade volumes, and we are interested in this overall effect.

To motivate our econometric specification of a two-part gravity model, we therefore consider a model of symmetric monopolistically competitive firms, as in Krugman (1980), thereby neglecting firm heterogeneity. In that model the extent of fixed bilateral market entry costs relative to operating profits in a market governs a firm's decision to serve the target market via exports or not.

[17] They are also important determinants of trade flows. In general, discrepancies on the levels of political competition, autocracy and durability reduce the probability of forming a PTA and the volume of trade. However, this fails for the difference in autocracy, which has a positive effect on PTA, and the difference in political competition which appears to increase exports.

A. *Theoretical Model*

Let us denote export-market specific fixed costs for firm b in country i to deliver goods to market j by $f_j(b)$. Each firm b supplies a single variety of the product and faces market-specific profits $\pi_j(b)$ in country $j = 1, \ldots, N$ of

$$\pi_j(b) = [\hat{p}_j(b) - \hat{z}_j(b)]\, c_j(b) - f_j(b). \tag{10}$$

In equation (10), $\hat{p}_j(b)$ denotes the consumer price of variant b and $\hat{z}_j(b)$ are the associated marginal costs of supplying variant b to consumers in j (including marginal production costs and trade costs). Unlike Helpman, Melitz, and Rubinstein (2008), let us assume for the model outset that all producers in country i are symmetric with respect to $\hat{z}_j(b)$ and $f_j(b)$.[18] As a consequence, we may drop product index b throughout our analysis and index products by their country of origin. Then, we may substitute $\pi_j(b) = \pi_{ij}$, $\hat{p}_j(b) = \hat{p}_{ij}$, $\hat{z}_j(b) = \hat{z}_{ij}$, $c_j(b) = c_{ij}$, and $f_j(b) = f_{ij}$ for all variants delivered by i-borne producers to consumers in j.

Firms in i will now maximize profits across all markets by setting identical mill prices p_i for consumers everywhere. With iceberg-type trade costs t_{ij} for exports from i to j, the relationship between consumer prices and mill prices is determined as $\hat{p}_{ij} = p_i t_{ij}$. Similarly, marginal delivery costs relate to marginal production costs by $\hat{z}_{ij} = z_i t_{ij}$, and shipments at the firm level may be defined as $x_{ij} \equiv c_{ij} t_{ij} = p_i^{-\sigma} t_{ij}^{1-\sigma} P_j^{\sigma-1} y_j$.

Accordingly, we may rewrite equation (10) as

$$\pi_{ij} = (p_i - z_i)x_{ij} - f_{ij}. \tag{11}$$

Notice that fixed entry costs f_{ij} are specific to an import market. Consequently, i-borne firms will decide to supply goods to consumers in j only if operating profits $(p_i - z_i)x_{ij}$ cover the market-specific fixed costs f_{ij}. With monopolistic competition, a constant elasticity of substitution σ between products, and a fixed markup over marginal production costs, operating profits per unit of output are $(p_i - z_i) = p_i/\sigma$ and i-borne firms will supply market j only if $p_i x_{ij} \geq \sigma f_{ij}$. Let us define an indicator function $\mathcal{I}_{ij}$ which is unity, if $p_i x_{ij} \geq \sigma f_{ij}$, and zero else. After defining the number of producers in country i as n_i, we may write aggregate nominal goods exports from i to j in equilibrium as

$$n_i p_i x_{ij} \equiv X_{ij} = \mathcal{I}_{ij} n_i p_i^{1-\sigma} t_{ij}^{1-\sigma} P_j^{\sigma-1} y_j. \tag{12}$$

As in Anderson and van Wincoop (2003), a country's world (including intranational) sales add up to GDP and we may state:

$$y_i = \left(n_i p_i^{1-\sigma}\right) \sum_{j=1}^{N} \left(\mathcal{I}_{ij} t_{ij}^{1-\sigma} P_j^{\sigma-1} y_j\right). \tag{13}$$

[18] We will estimate some models below that accommodate the case of heterogeneous firms. This can be done by introducing a polynomial control function which is based on the linear predictions in the extensive margin model, similar to Helpman, Melitz, and Rubinstein (2008). See below for further details.

Now, after defining $y_W = \sum_{i=1}^{N} y_i$, we may substitute $(n_i p_i^{1-\sigma})$ by $y_i / y_W \Pi_i^{1-\sigma}$ in (12) to obtain an equivalent expression for nominal aggregate bilateral exports to the one in equation (1). Yet, unlike in (1), zero bilateral exports may surface in the non-stochastic part of the model:

$$X_{ij} = \mathcal{I}_{ij} \frac{y_i y_j}{y_W} t_{ij}^{1-\sigma} \, \Pi_i^{\sigma-1} \, P_j^{\sigma-1}. \tag{14}$$

Analogous to the discussion in Section I, the unobserved $\Pi_i^{1-\sigma}$ and $P_j^{1-\sigma}$ can be computed as implicit solutions to the system of $2N$ equations

$$\Pi_i^{1-\sigma} = \sum_{j=1}^{N} \left(\mathcal{I}_{ij} t_{ij}^{1-\sigma} P_j^{\sigma-1} y_j / y_W\right); \qquad P_j^{1-\sigma} = \sum_{i=1}^{N} \left(\mathcal{I}_{ij} t_{ij}^{1-\sigma} \Pi_i^{\sigma-1} y_i / y_W\right), \tag{15}$$

where $\Pi_i^{1-\sigma}$ and $P_j^{1-\sigma}$ are the equivalent expressions to the ones in equation (2), but allowing for zero trade flows.

B. *An Empirical Two-Part Model of Trade*

We consider now estimation of a stochastic version of the gravity model with zero trade flows as in (14):

$$X_{ij} = \mathcal{I}_{ij} \exp(\mathbf{Z}_{ij}'\beta + \delta PTA_{ij} + \alpha_i + \gamma_j)\epsilon_{ij}. \tag{16}$$

Taking expectations and using the law of iterated expectations we can write the CEF as

$$\begin{aligned} E(X_{ij} | \cdot) &= \Pr(\mathcal{I}_{ij} = 1 | \cdot) E(\exp(\mathbf{Z}_{ij}'\beta + \delta PTA_{ij} + \alpha_i + \gamma_j)\epsilon_{ij} | \cdot, \mathcal{I}_{ij} = 1) \\ &= \Pr(\mathcal{I}_{ij} = 1 | \cdot) E(X_{ij} | \cdot, \mathcal{I}_{ij} = 1). \end{aligned} \tag{17}$$

This is a two-part model which allows to decompose the effects of the explanatory variables on exports into an effect on the extensive country margin—i.e., the decision to export to a country at all—and on the intensive margin—i.e., on the value of exports conditional on positive exports. In the baseline model (8), the estimated effect represents some average of these two. Two-part econometric models have been discussed in econometrics for some time (John G. Cragg 1971; Naihua Duan et al. 1984), but have not been implemented in the empirical trade literature so far, to the best of our knowledge.

To complete the specification of the two-part model and make it operational, functional forms for the probability of trading and the expected trading volume have to be defined. Retaining endogeneity of PTA in exports, we postulate for the second part of (17) a similar relationship as the one used before,

$$E(X_{ij} | \mathbf{Z}_{ij}, \mathbf{W}_{ij}, PTA_{ij}, \mathcal{I}_{ij} = 1) = \lambda_{ij} \, \Psi_{ij}, \tag{18}$$

where λ_{ij} and Ψ_{ij} are analogous to the expressions in (9). However, note that as this functional form is now assumed to hold for positive exporters only, and not for all observations as in (8)–(9), the parameters β, δ and ϑ in (18) do not denote the same quantities as in the model of Section IV.

Let us now turn to the first part of the model, the probability of country i to serve country j via exports at all. For this purpose, the model for $\mathcal{I}_{ij}$ as defined by equation (11) is translated into a stochastic process

$$\mathcal{I}_{ij} = \begin{cases} 1 & \text{if} \quad \mathbf{Q}_{ij}'\omega + \kappa PTA_{ij} \geq \xi_{ij}, \\ 0 & \text{else,} \end{cases} \tag{19}$$

where the vector $\mathbf{Q}_{ij}$ is a set of observable variables determining positive exports (i.e., positive profits for firms in i which are specific to market j), ω are the corresponding unknown parameters, κ is the parameter of the PTA indicator variable, and ξ_{ij} is a stochastic term. Note that $\mathbf{Q}_{ij}$ may but need not contain the same elements as $\mathbf{Z}_{ij}$. Since PTA membership is an endogenous determinant of the positive value of exports, it would be awkward to assume that it is exogenous to the decision to export at all from i to j. Therefore, we explicitly allow for dependence between ξ_{ij} and PTA_{ij}. With a binary dependent variable ($\mathcal{I}_{ij}$) and a binary endogenous regressor (PTA_{ij}) at hand, we follow a large literature in modeling the two binary processes by means of a bivariate probit model (cf. Chiara Monfardini and Rosalba Radice 2008, for some recent applications). Then, the probability of trading conditional on PTA membership can be written as (see, e.g., William H. Greene 2008)

$$\Pr(\mathcal{I}_{ij} = 1 \,|\, \mathbf{Q}_{ij}, \mathbf{W}_{ij}, PTA_{ij}) = \frac{\Phi_2[(2PTA_{ij} - 1)\mathbf{W}_{ij}'\theta, \mathbf{Q}_{ij}'\omega + \kappa PTA_{ij}, (2PTA_{ij} - 1)\rho_{v\xi}]}{\Phi[(2PTA_{ij} - 1)\mathbf{W}_{ij}'\theta]}, \tag{20}$$

where Φ_2 denotes the bivariate normal cumulative distribution function and $\rho_{v\xi}$ the correlation between v and ξ.

Thus, the impact of a variable on the CEF (17) is modeled in a very flexible manner in the two-part model, allowing a variable to have different effects in each part of the two components of (17). For instance, it is possible for a variable to have a strong impact on the extensive country margin—the probability of initiating exports to a given country which is determined mainly by ω and κ at given $\mathbf{Q}_{ij}$ and PTA_{ij}—but to have a small impact on the intensive margin—an increase of the value of positive bilateral exports resulting principally from β at given regressors.

A convenience of such a model is that the two parts, (18) and (20), can be estimated separately. Thus, consistent estimates of the parameters of (20), ω, κ, θ as well as the degree of endogeneity of PTA (as measured by the correlation between PTA and ξ_{ij}) can be obtained by standard maximum likelihood estimation. As for the parameters from (18), we can use the same two-stage PML procedures described in Section IV, and include only the observations with positive exports in the estimation.

C. *A Two-Part Model with Correlated Errors and Firm Heterogeneity*

The two-part model presented in the previous sub-section differs from alternative approaches suggested in the recent literature to discriminate between effects at the extensive and intensive country margins of trade. The dominant procedure in the current literature is to estimate some form of sample selection model (see Helpman, Melitz, and Rubinstein 2008; and Santos Silva and Tenreyro 2008), most often a log-linear Heckit model.

While it is generally acknowledged that the statistical properties and robustness of sample selection models are often inferior to the ones of the two-part model (Duan, et al. 1984; Manning, Duan, and William H. Rogers 1987; Siu Fai Leung and Shihti Yu 1996; William H. Dow and Edward C. Norton 2003), proponents of sample selection models have advocated the use of the former over the latter based on its explicit modeling of the correlation structure between the errors in the two equations.[19] The two-part model does not estimate any covariance terms between two error vectors. However, it is consistent under some general classes of joint distributions which allow for stochastic dependence between error terms. In particular, the two-part model allows for joint distributions of the errors which are excluded by assumption under sample selection models (see Duan, et al. 1984). If the error terms are independent across the participation and outcome equations, the two-part model is efficient relative to the sample selection model. If the errors are not independent, it is difficult to decide between two-part versus selection models as they are not nested.

We decided to address correlated disturbances between the intensive and extensive margin by estimating an additional two-part model with a semiparametric control-function. We follow Helpman, Melitz, and Rubinstein (2008) to postulate that the disturbances of the two models are independent conditional on some nonlinear function of the predictions of the extensive margin equation. In the Heckit approach taken by Helpman, Melitz, and Rubinstein (2008), the linear prediction of the extensive margin model enters in a nonlinear form by inclusion of the inverse Mills' ratio, i.e., the ratio of the probability density function and the cumulative normal distribution function evaluated at the linear prediction of the participation equation. We denote this linear prediction by

$$\hat{\eta}_{ij} \equiv \mathbf{Q}'_{ij}\hat{\omega} + \hat{\kappa} PTA_{ij}.$$

In our application, the inverse Mills' ratio would not contribute significantly to the explanatory power since it is almost linear over a large part of the relevant range of its argument.[20]

[19] The *participation* equation which, in our case, corresponds to the extensive country margin model of trade, and the *outcome* equation which corresponds to the intensive country margin model.

[20] This is true for the models we report below as well as ones that use religion as in Helpman, Melitz, and Rubinstein (2008) as an identifying instrument in the participation equation for the extensive country margin of exports. The coefficient estimates are very robust to the inclusion of religion as an identifying instrument for the extensive country margin of exports. However, we report results which exclude this variable for the sake of keeping an additional 506 observations in the sample for which the religion variable is not available.

Thus, we model the correlation among disturbances conditional on a nonlinear function of $\hat{\eta}_{ij}$ by including polynomial terms up to a fourth order (excluding the linear term for reasons of collinearity). Not only is this approach more flexible than the inclusion of the Mills' ratio alone, but it also has the added advantage that the polynomial control function captures possible nonlinearities due to the firm extensive margin. Such additional nonlinearities are present if firms are in fact heterogenous. Notice that either problem is addressed by a (slightly different) nonlinear function about $\hat{\eta}_{ij}$ also in Helpman, Melitz, and Rubinstein (2008). We argue that the fourth-order polynomial in $\hat{\eta}_{ij}$ is a flexible approximation of both nonlinear functions. While within this semiparametric approach it is impossible to distinguish between error correlation and firm heterogeneity, this is irrelevant for our purpose of analyzing country margins of trade.[21] The (combined) presence of error correlation and firm heterogeneity can be tested for by an F-test of joint significance on the coefficients of the approximating polynomial function.

D. *Estimation Results*

Here we discuss the parameter estimates from the models described in Section VB and Section VC. Similar to Table 2, Table 3 summarizes the estimates of two alternative models of nominal bilateral exports in US dollars (X_{ij}). Again, every column gives parameters and standard errors of covariates of interest in the pertinent equations. Yet, now we distinguish between the process generating zero versus positive exports on the one hand, and the process generating alternative positive values of exports on the other. The former is captured by a probit model for $\mathcal{I}_{ij}$ as explained in Section VB, while the latter is estimated via Poisson PML.

Columns 2 to 4 give parameters and standard errors when treating PTA_{ij} as exogenous. Both column 3 and 4 represent trade volume equations. The difference between those two is that the model in column 4 controls for higher-order polynomial terms in $\hat{\eta}_{ij}$ while the one in (3) does not. Hence, the specification in (4) implicitly accommodates correlation between the disturbances of the extensive country margin equation in column 2 and the intensive country margin model in column 3 as well as heterogeneous firms akin to the approach of Helpman, Melitz, and Rubinstein (2008). In columns 6 to 8, we treat PTA_{ij} as endogenous. To model this, in addition to the two equations for $\mathcal{I}_{ij}$ and exports, a third equation for PTA is needed. Endogeneity of PTA_{ij} in $\mathcal{I}_{ij}$ is captured by a recursive bivariate probit model as summarized in columns 5 and 6. While we use the same specification for the latent process behind the extensive country margin $\mathcal{I}_{ij}$ (i.e., the process η_{ij}) as for positive exports $\left(\text{or } E(X_{ij} \mid X_{ij} > 0)\right)$, the results are virtually identical to a model where we use religion as an additional regressor in the extensive margin model as in Helpman, Melitz, and Rubinstein (2008). Similar to columns 3 and 4 with exogenous PTAs, we estimate models (7) and (8) with endogenous PTAs where the difference lies in the polynomial function about $\hat{\eta}_{ij}$ which is present in column 8 but not in 7. The coefficients of the polynomial function are jointly significant (in both

[21] This is in contrast to work analyzing firm margins of trade, for which the distinction between firm extensive margin and error correlation is crucial (Helpman, Melitz, and Rubinstein 2008).

TABLE 3—ESTIMATION RESULTS FOR TWO-PART GRAVITY MODELS FOR TRADE

	Exogenous PTA			Endogenous PTA			
Regression Estimator (1)	$\Pr(X > 0)$ Probit (2)	$E(X \mid X > 0)$ Poisson (3)	$E(X \mid X > 0)$ Poisson (4)	$\Pr(\text{PTA} = 1)$ Biv.Probit (5)	$\Pr(X > 0)$ Biv.Probit (6)	$E(X \mid X > 0)$ Poisson (7)	$E(X \mid X > 0)$ Poisson (8)
PTA	0.3515	0.3698	0.3642	—	−0.0307	1.2118	1.2701
	(0.0559)	(0.0711)	(0.0706)		(0.1326)	(0.3715)	(0.3961)
DIST	−1.1454	−0.6209	−0.6049	−1.0689	−0.9448	−0.3873	−0.4958
	(0.0466)	(0.0344)	(0.0395)	(0.0388)	(0.0434)	(0.0670)	(0.1135)
BORD	−0.4276	0.6478	0.6948	−0.4469	−0.0647	0.7566	0.8124
	(0.1848)	(0.0596)	(0.0603)	(0.1184)	(0.1673)	(0.0759)	(0.0812)
LANG	0.6341	0.2108	0.1907	−0.0974	0.6258	0.2241	0.2368
	(0.0628)	(0.0627)	(0.0613)	(0.0616)	(0.0616)	0.0655)	0.0922)
CONT	0.1041	0.2897	0.2926	0.7794	0.3095	0.0930	−0.0797
	(0.0596)	(0.0656)	(0.0642)	(0.0479)	(0.0581)	(0.1568)	(0.1531)
DURAB	−0.0070	−0.0028	−0.0026	−0.0071	−0.0018	−0.0043	−0.0040
	(0.0016)	(0.0006)	(0.0006)	(0.0010)	(0.0014)	(0.0009)	(0.0009)
POLCOMP	−0.0027	0.0804	0.0918	−0.0483	0.0042	0.0947	0.0920
	(0.0087)	(0.0198)	(0.0201)	(0.0089)	(0.0086)	(0.0262)	(0.0269)
AUTOC	0.0011	−0.0594	−0.0681	0.0479	−0.0069	−0.1296	−0.1283
	(0.0093)	(0.0301)	(0.0298)	(0.0097)	(0.0092)	(0.0305)	(0.0317)
CURCOL	−0.1339	0.3802	0.3685	0.5218	−0.1638	0.5391	0.4842
	(0.3752)	(0.1872)	(0.1807)	(0.2377)	(0.3396)	(0.1784)	(0.1788)
$(\hat{\eta}_{ij})^2$	—	—	−0.0377	—	—	—	−0.1129
			(0.0044)				(0.0483)
$(\hat{\eta}_{ij})^3$	—	—	0.0094	—	—	—	0.0359
			(0.0012)				(0.0365)
$(\hat{\eta}_{ij})^4$	—	—	−0.0006	—	—	—	−0.0053
			(0.0001)				(0.0063)
COLONY	—	—	—	0.1318	—	—	—
				(0.1899)			
COMCOL	—	—	—	0.4505	—	—	—
				(0.0757)			
SMCTRY	—	—	—	1.2244	—	—	—
				(0.2210)			
$\hat{\rho}_{v\xi}$	—	—	—	—	0.2895	—	—
					(0.0834)		
$\hat{\vartheta}$	—	—	—	—	—	−0.4016	−0.4373
						(0.1710)	(0.1770)
F-stat.	—	—		48.4092	3.7782	1.6329	5.5857
p-value of *F*	—	—		0.0000	0.2864	0.6519	0.1336
F-stat. OIR	—	—		—	2.6605	1.4530	3.0029
p-value of OIR	—	—		—	0.1029	0.4836	0.2228
Number	13,500	9,891	9,891	15,750	15,750	9,891	9,891
Countries	126	126	126	126	126	126	126

Notes: All regressions include importer and exporter fixed effects. In the second column, the number of observations is reduced due to countries that export to the whole "world" and which are dropped from the estimation. These are Belgium, Canada, Switzerland, China, Germany, Denmark, Finland, France, GB, Indonesia, Italy, Japan, Korea, Netherlands, Norway, Sweden and USA. The sources of the data are the United Nations' World Trade Database and the Centre d'Etudes Prospectives et d'Informations Internationales. "*F*-stat." and "*p*-value of *F*" refer to a test of joint significance of *COLONY*, *COMCOL* and *SMCTRY* in the respective equation. The *F*-statistic should be significantly different from zero in column 5 but not in 6–8. "*F*-stat. OIR" and "*p*-value of OIR" refer to a test for overidentifying restrictions in the corresponding log-linear IV model.

Table 4—Partial (i.e., Non-General-Equilibrium) Effects of PTA

Model (1)	Total (2)	Decomposition: Ext. marg. (3)	Decomposition: Int. marg. (4)
Exogenous PTA			
Single index Table 2, column 3	74.15	—	—
Two-part Table 3, columns 2, 3	67.92	19.59	48.33
Two-part with $\hat{\eta}_{ij}$-terms Table 3, columns 2, 4	66.98	19.52	47.45
Endogenous PTA			
Single index Table 2, columns 4, 5	214.90	—	—
Two-part Table 3, columns 5–7	235.02	−0.60	235.62
Two-part with $\hat{\eta}_{ij}$-terms Table 3, columns 5, 6, 8	255.12	−0.63	255.75

Notes: Ext. marg. and int. marg. are the partial effects of PTA membership on the extensive and intensive margins of exports, respectively. Partial effects for two-part models are evaluated at the average of explanatory variables.

the exogenous case in column 4 as well as in column 8 where PTA is endogenous) suggesting some role for firm heterogeneity and/or error correlation between the decision to trade and trade flows.[22]

Consider the model imposing exogeneity first. Every variable has two associated parameters, one corresponding to the extensive margin in column 2 and one for the intensive margin of trade in column 3 or column 4. Almost all coefficients have the same sign in the extensive versus the intensive margin models. Exceptions are *BORD*, *CURCOL*, and *POLCOMP*—the latter being insignificant in column 2—, which appear to be impediments to start trading but foster trade flows once exports are positive. The result about *BORD* and *CURCOL* is in line with previous research (see Helpman, Melitz, and Rubinstein 2008).

To give an impression of the implied effect on trade flows, Table 4 shows the total partial effect of PTA on trade. By partial effect we mean here the effect that PTA has on exports keeping everything else constant, including the general equilibrium effects which work through the change in the multilateral resistance terms: $E(X\,|\,PTA = 1, \mathbf{Z})/E(X\,|\,PTA = 0, \mathbf{Z}) - 1$. A quantification by means of counterfactual analysis, taking into account third-country effects present in the MR terms in (15) and GDP through equation (13) is reported in Section VI.

With an estimated coefficient of $\hat{\delta} = 0.55$, the partial effect in the single index model from Table 2 is 74 percent $(= \exp(\hat{\delta}) - 1)$, which is reasonably

[22] We also compared the models based on goodness-of-fit (not reported). The estimated mean square error (MSE) of the two-part model with $\hat{\eta}$-terms is about 15 percent lower than the other two-part model's. In turn, this two-part model has a MSE which is substantially lower than that of a conventional log-linear Heckit or parametric Helpman, Melitz, and Rubinstein (2008) estimator.

close to the total partial effect at the average in the two-part model, 68 percent $(= [\Phi(\mathbf{Q}_{ij}'\hat{\omega} + \hat{\kappa})/\Phi(\mathbf{Q}_{ij}{}'\hat{\omega})]\exp(\hat{\delta}) - 1$, where sample averages are used for the variables in $\mathbf{Q}_{ij}$), so that we may say that they indeed measure the same quantity. The two-part model allows us to decompose this total effect into the contributions from the extensive and intensive margin:

$$\frac{E(X \mid \text{PTA} = 1, \mathbf{Z}) - E(X \mid \text{PTA} = 0, \mathbf{Z})}{E(X \mid \text{PTA} = 0, \mathbf{Z})}$$

$$= \frac{\Phi(\mathbf{Q}'\omega + \kappa) - \Phi(\mathbf{Q}'\omega)}{\Phi(\mathbf{Q}'\omega)} \frac{(\exp(\delta) + 1)}{2}$$

$$+ [\exp(\delta) - 1] \frac{\Phi(\mathbf{Q}'\omega + \kappa) + \Phi(\mathbf{Q}'\omega)}{\Phi(\mathbf{Q}'\omega)},$$

where the first term is the partial effect at the extensive margin and the second term the one at the intensive margin. Evaluated at the average of the explanatory variables, these are estimated to be 20 percentage points and 48 percentage points, respectively, in our sample (i.e., around 70 percent of the partial effect is found to be attributable to the intensive country margin).

Let us contrast this finding with the results obtained when letting PTA be potentially endogenous. We use the same variables here as before—*COLONY*, *COMCOL*, and *SMCTRY*—as identifying instruments for PTA in the two equations for $\mathcal{I}_{ij}$ (column 6 in Table 3) and positive exports (columns 7 and 8). As the distributional assumption identifies the bivariate probit model, we can test the three overidentifying restrictions (OIR) with an F-test. The null hypothesis that the restrictions hold cannot be rejected in neither part of the model (row "p-value of F" in columns 6, 7 and 8). The instruments also pass the OIR test in log-linearized instrumental variable models (row "p-value of OIR" in columns 6, 7, and 8)[23].

Let us now discuss the results about the error correlations and the coefficient of PTA. The estimate of ϑ is negative and significant, which is in line with our findings in Table 2. Hence, as before, selection into PTAs on unobservables is negative. A significant $\hat{\rho}_{v\xi}$ likewise suggests that there is endogeneity in the selection into exports decision. Here, we find evidence of positive self-selection based on unobservables, which is reflected in the overestimation of the impact of PTA on the decision to export at all when neglecting endogeneity of PTA. As a matter of fact, the results in Table 3 suggest that after controlling for endogeneity, PTA membership has an impact on the intensive margin, but does not significantly affect the extensive margin of trade, i.e., the country intensive margin accounts for the whole partial effect of PTA (see also the corresponding estimated partial effects in Table 4).

Such a result could, for instance, be explained by sufficiently high market-specific fixed entry costs which are unaffected by PTA formation, whereas marginal delivery costs are lowered by PTA membership. Note that these results are compatible with

[23] The test has two degrees of freedom here as one instrument is needed for identification.

empirical research at a disaggregate level which emphasizes PTA to be important in shaping some extensive margins of trade. Timothy J. Kehoe and Kim J. Ruhl (2009), for instance, found NAFTA to affect the extensive margin at the level (i.e., the number) of products traded. Extensive and intensive product margins are partly subsumed by, but not identical to, the country intensive margin.[24] Finally, the fact that the estimated correlations are of different signs is perfectly compatible with the general specification of the model. The differently signed correlations suggest that, after controlling for economic and political determinants, extensive and intensive margins of export appear to be driven by heterogeneous factors.

The remaining variables are only marginally affected by the change from the model for exogenous PTA to the one where it is endogenous, with the sole exception here being again *BORD*, which looses its significant negative impact on the extensive margin.

VI. Quantification and Discussion

We will illustrate the importance of considering both self-selection into PTAs and zero export flows by means of counterfactual analysis. In particular, we will compute the impact of PTA membership as observed in the year 2005 to a situation without any PTA membership in the same year, using a variety of different estimators and taking into account general equilibrium effects addressed in Sections I and VA.

The literature on the impact of endogenous PTA formation on trade suggests a positive parameter estimate on nominal bilateral exports. For instance, Baier and Bergstrand (2009) report estimates of average treatment effects of in between 0.68 (using the matching estimator of Alberto Abadie and Guido W. Imbens (2006), for the year 2000; implying an effect of about 97 percent) and 2.36 (using the same approach for the year 1990). While these estimates lie in a similar range as the ones reported in previous work and take non-linear effects of trade costs as possible determinants of PTA formation into account, they do not consider non-linear general equilibrium effects of PTAs on exports. Baier and Bergstrand (2007) acknowledge general equilibrium effects with panel data but assume that PTA membership is exogenous. However, the average treatment effects from their preferred models are still very close to the cross-sectional endogenous treatment effects in their more recent paper, amounting to 0.62 (implying an effect of about 86 percent) and 0.54 (implying an effect of about 72 percent). Relative to Baier and Bergstrand's (2009), Magee's (2003) estimated PTA-effects on trade seem rather large: they lie between 300 percent and 800 percent. However, these estimates do not account for fixed country effects in both trade volume and PTA equation.

Unlike previous work, our quantification of PTA effects on trade flows respects general equilibrium effects, accounts for the differential impact of PTAs on the extensive and intensive margins of exports, and treats PTAs endogenously. Finally,

[24] A change at the extensive product margin as in Feenstra and Hiau Looi Kee (2008) or Kehoe and Ruhl (2009) would be equivalent to a change at the extensive country margin only for the first product(s) traded bilaterally. Otherwise, an expansion at the extensive product margin for any pair of countries is measured by an increase in the intensive country margin of exports in an aggregate analysis as ours.

we will also infer the importance of something that did not surface in the debate about PTA effects on trade yet: that most-favored nation tariffs are heterogeneous so that PTA membership does not bring about identical tariff reductions across country-pairs (see Anderson and van Wincoop 2002, for a treatment of tariff effects in their general equilibrium model).

Starting point of the quantification are the parameter estimates summarized in Tables 2 and 3. Note that so far we did not need to rely on any specific underlying model. Our estimation equations leading to the econometric specification for the parameter estimation are perfectly consistent with a wide range of recent international trade models.[25] Specifically, it captures new trade theory models with love-of-variety preferences and homogeneous firms à la Krugman (1980), the Anderson and van Wincoop (2003) exchange economy, the Helpman, Melitz, and Rubinstein (2008) model allowing for firm heterogeneity and zero trade flows, the Eaton and Kortum (2002) Ricardian model, and the Melitz and Gianmarco I. P. Ottaviano (2008) model with quasi-linear quadratic preferences and endogenous mark-ups. However, if one wants to go further and run a counterfactual analysis, it is necessary to adopt one specific model and use the implied structural model equations. In the subsequent analysis, we apply the Anderson and van Wincoop (2003) framework. Hence, in addition to the parameter estimates, we use the assumption that countries are endowment economies and exports are related to exporter and importer GDP as well as multilateral resistance terms as in equation (1).

For a quantification of the general equilibrium-consistent average treatment effect of observed PTA membership on exports, we need to determine counterfactual bilateral exports in the absence of PTA membership. For this, we set the binary PTA indicator to zero and solve the system of $2N$ equations of exporter and importer MR terms in (15).[26] This can be done by assuming that PTA membership is associated with heterogeneous tariff reductions or not.[27] Irrespective of whether heterogeneous tariffs are acknowledged or not, PTA formation has an impact on GDP and the latter has to be considered in the solution of (15) and in the outcome equation for the intensive margin of exports, i.e., in (14). Tables 5 and 6 summarize the predicted effects of PTA formation on trade among PTA members relative to non-members for the models estimated in Tables 2 and 3, respectively.[28] For each model,

[25] This is due to the fact that the differences between models pertain to country-specific variables, which are captured by country-fixed effects in our estimation equations.

[26] Since σ, the elasticity of substitution between products, is not known, its level has to be estimated or to be assumed. In the model of Anderson and van Wincoop (2003), it can not be estimated since the model does not impose enough structure. So, we follow them in setting $\sigma = 5$. Note that they find trade predictions to be fairly insensitive to the choice of different values of σ. (If one used a model which specified the supply side explicitly—such as a multi-country version of Krugman's (1980) model—one would be able to estimate σ.)

[27] To account for heterogeneous tariffs, we use data on tariff revenues in total trade flows from the World Bank's World Development Indicators 2007, assume that tariff rates are identical vis-à-vis all PTA nonmembers, and apply these tariffs to trade flows of all trading partners of a country in the counterfactual abolishment of preferential trade liberalization. In principal, one could replace PTA_{ij} by an appropriately defined (endogenous) tariff variable and apply the framework suggested here. However, tariffs may be inaccurately measured and PTA membership may entail more than just a bilateral reduction in tariffs. Therefore, we prefer approximating tariff effects as indicated but employ the binary indicator variable in the regressions.

[28] According to Walras' law, absolute trade effects are impossible to gauge in general but they have to be expressed relative to one country-pair or relative to a group of pairs.

Table 5—Counterfactual Results with Homogeneous Tariff Rates

	One-part models (Table 2)		Two-part models[b] (Table 3)			
	Exog. PTA	Endog. PTA	Exog. PTA		Endog. PTA	
Based on columns in respective table	(3)	(4)–(5)	(2)–(3)	(2), (4)	(5)–(7)	(5), (6), (8)
Average percentage increase of trade flows of PTA members relative to non-members	58.97	161.36	41.80	45.75	219.36	234.57
ΔX_{ij} among PTA members in percent:[a]						
mean	39.3965	101.6426	14.2631	12.6884	59.4102	61.8560
std. dev.	29.5054	94.7278	15.1128	14.5677	68.1883	73.9056
min	−10.9748	−22.2771	−25.9965	−25.1213	−63.4330	−64.9363
max	141.6656	514.3927	62.0881	61.9867	381.9351	431.8087
Number of PTA member pairs with						
positive effect	3,456	3,450	2,749	2,719	2,701	2,674
negative effect	50	56	229	259	277	304
ΔX_{ij} among PTA non-members in percent:[a]						
mean	−4.8554	−9.0345	−6.6081	−6.7792	−18.8946	−19.5906
std. dev.	16.1757	32.7673	8.8012	9.0138	22.7672	23.7805
min	−41.8862	−67.5594	−45.9182	−45.3577	−87.3477	−88.5390
max	38.6748	95.6077	10.4128	13.0945	27.7728	32.5374
Number of PTA non-member pairs with						
positive effect	4,108	3,916	691	683	473	498
negative effect	8,136	8,328	6,222	6,230	6,440	6,415

Notes: There are 3,632 PTA member pairs, whereof in 654 $\Delta X_{ij} = 0$ occurred. Of the 12,244 PTA non-member pairs, in 5,331 cases the model predicted $\Delta X_{ij} = 0$.
[a] ΔX_{ij}: base scenario trade flows minus counterfactual trade flows relative to counterfactual trade flows in percent; std. dev.: standard deviation.
[b] In the two-part models, ΔX_{ij} was calculated in the sub-sample of pairs with positive trade flows X_{ij} in both the benchmark and the counterfactual equilibrium.

we distinguish between effects that assume that PTAs alter homogeneous tariffs (in Table 5) preferentially versus ones that alter heterogeneous tariffs (in Table 6).

In a nutshell, the figures in the tables suggest the following conclusions. First, trade among PTA members increases due to preferential tariff abolition. For instance, the model which assumes exogenous PTA formation, no specific process for the extensive margin to export, and no heterogeneous tariff effect on trade and GDP points to an increase in nominal exports among PTA members relative to nonmembers by about 59 percent relative to an equilibrium without any PTAs. This is reflected in the number which is given in the outer left column of the top row of Table 5 labeled "*Average percentage increase of trade flows of members relative to non-members*." The estimated effect is about 102 percentage points higher with endogenous PTA formation (about 161 percent higher exports among PTA members relative to nonmembers than without PTAs; see the results in the column labeled (3)–(4) at the top of Table 5). Ignoring the heterogeneity of tariffs brings about a negligible bias in our application.[29] To see the latter, compare the results at the top of Table 5 with the corresponding ones in Table 6.

[29] In all models, the predicted effect of PTA membership on members' versus nonmembers' bilateral trade is slightly lower when considering heterogeneity of most-favored nation tariffs. The relatively small bias from

Table 6—Counterfactual Results with Heterogeneous Tariff Rates

	One-part models (Table 2)		Two-part models[b] (Table 3)			
	Exog. PTA	Endog. PTA	Exog. PTA		Endog. PTA	
Based on columns in respective table	(3)	(4)–(5)	(2)–(3)	(2), (4)	(5)–(7)	(5), (6), (8)
Average percentage increase of trade flows of PTA members relative to non-members	58.61	160.88	41.68	45.69	219.13	234.48
ΔX_{ij} among PTA members in percent:[a]						
mean	40.7396	103.1379	13.7920	12.5047	58.7820	61.3718
std. dev.	30.5686	96.0032	15.1707	14.8975	67.8647	73.5924
min	−10.0238	−21.7366	−27.4492	−26.8490	−64.0591	−65.5265
max	142.8524	516.1415	61.1247	62.5190	376.4445	425.3974
Number of PTA member pairs with						
positive effect	3,458	3,448	2,736	2,695	2,700	2,674
negative effect	48	58	242	283	278	304
ΔX_{ij} among PTA non-members in percent:[a]						
mean	−3.7172	−8.2642	−6.8230	−6.7841	−19.0597	−19.7246
std. dev.	16.5521	33.0922	8.9757	9.1057	22.8397	23.8268
min	−42.0441	−67.6475	−46.4574	−45.8900	−87.4854	−88.6601
max	39.7881	96.4289	10.1447	12.6403	26.9016	32.1130
Number of PTA non-member pairs with						
positive effect	4,484	4,008	658	723	458	483
negative effect	7,760	8,236	6,255	6,190	6,455	6,430

Notes: There are 3,632 PTA member pairs, whereof in 654 $\Delta X_{ij} = 0$ occurred. Of the 12,244 PTA non-member pairs, in 5,331 cases the model predicted $\Delta X_{ij} = 0$.

[a] ΔX_{ij}: base scenario trade flows minus counterfactual trade flows relative to counterfactual trade flows in percent; std. dev.: standard deviation.

[b] In the two-part models, ΔX_{ij} was calculated in the sub-sample of pairs with positive trade flows X_{ij} in both the benchmark and the counterfactual equilibrium.

Modeling the process of endogenous selection into positive exports separate from the non-linear process of positive exports is relatively important. It raises the predicted effect of PTA formation with endogenous PTAs on insiders' trade relative to other country-pairs. In the column labeled (5)–(7) of Table 5, the estimated comparative static effect of PTA membership amounts to 219 percent, which exceeds the one-part model-based result of 161 percent by about 58 percentage points. In the two-part model, which allows for correlated disturbances between the extensive and intensive country margin equations at the outer right of Table 5, the corresponding estimated comparative static effect is about 235 percent. Of the latter effect on exports, the lion's share is contributed by the intensive margin.[30] Overall, the estimated (long-run) effects of PTA membership on bilateral trade are quite large.

ignoring tariff heterogeneity has to do with the fact that, on average, most-favored nation tariffs are relatively homogeneous across countries in 2005 so that capturing tariff effects by a binary PTA indicator variable does not conceal much information.

[30] Of 9,891 country-pairs with predicted positive bilateral exports in the cum-PTA benchmark equilibrium, 69 are predicted to stop exporting if all PTAs were abandoned when using the exogenous-PTA probit model in column 2 of Table 3. With the endogenous-PTA model in column 6 of Table 3, 106 are predicted to stop exporting in counterfactual general equilibrium relative to the benchmark equilibrium. This result is based on estimates which disregard the fact that (most-favored nation) tariffs are heterogeneous across countries so that preferential

Moreover, Tables 5 and 6 indicate that a focus on PTA effects on *average* trade flows—as had been done in most of the previous work on endogenous PTA effects on trade flows—conceals the sizable variation effects across country pairs.[31] To see this, consider the two blocs of results in the lower parts of Tables 5 and 6. There, we report four moments of the distribution of the percentage changes of bilateral exports both of PTA members (at the center of each table) as well as of non-members (at the very bottom of each table): the mean, the standard deviation, the minimum, and the maximum effect for each model.[32] Obviously, most of the models display a standard deviation of effects within the groups of PTA members and non-members, which exceeds the average effect. The variation in the effects is entirely due to the relevance of heterogeneity across countries in general equilibrium. Hence, the underlying theoretical model suggests that the treatment effect of PTA membership is inherently heterogeneous. The results even point to negative effects from the simultaneous implementation of PTAs in the world economy on some PTA members (accruing to third country effects of foreign PTAs). Similarly, there are even PTA non-members which gain from the simultaneous implementation of foreign PTAs. PTA members face positive and PTA non-members negative effects of PTA formation on trade flows only on *average*.

Altogether, these findings suggest that the empirical models proposed here may help to estimate effects of endogenous PTA effects on trade flows which have appeal from both a theoretical and an empirical perspective. First, proposed models principally allow for a disproportionate number of zero trade flows *and* endogenous PTA membership which previously proposed estimators for gravity models did not allow for (and accommodated only one or the other). Second, the proposed models allow for estimation of effects which fully account for general equilibrium effects of PTA membership associated with GDP responses to membership and ultimately heterogeneous treatment effects of PTA formation. For instance, recently proposed microeconometric methods (such as propensity score matching or switching regression) did not share this feature.

VII. Concluding Remarks

This paper proposes non-linear econometric techniques for the analysis of trade policy effects on bilateral trade flows which subsume three features: they pay specific attention to zeros in bilateral trade matrices; they allow trade policy variables—such as binary preferential trade agreement (PTA) indicators but eventually also

trade liberalization is associated with tariff reductions of different magnitude across country-pairs. In order to disentangle PTA-induced effects on exports that arise through changes at the extensive and intensive margins of trade we proceeded as follows. First, we calculated the total effect on trade by using estimates of the two-part model with endogenous PTAs—including effects on the extensive and intensive margins. Then, we calculated an alternative counterfactual by holding the margin constant at the benchmark equilibrium. The latter, leads to results that are very similar to the ones for the one-part models, where endogenous selection into positive exports is not accounted for.

[31] That treatment effects tend to be heterogeneous across the treated is widely acknowledged in other fields of economics (see Marianne P. Bitler, Jonah Gelbach, and Hilary W. Hoynes 2006, for an example in public economics). However, empirical international economists tend to focus on average effects of treatments such as PTA membership and other treatments on outcome of interest and tend to ignore that theoretical models often would suggest heterogeneous treatment effects.

[32] Notice that, for all two-part models, we provide these figures only for the subsample of country-pairs with positive exports.

continuous trade policy measures—to be endogenous; and they account for nonlinear effects of trade policy and trade costs in stylized general equilibrium models. All of these features have been judged as being important in recent empirical work in international economics, but no attempt has been made to address them in a unified framework as we do.

Apart from addressing the issue from an econometric perspective and from summarizing methodical frameworks for empirical work on the matter, we apply the suggested procedures to estimate general equilibrium-consistent effects of PTA membership on bilateral trade flows in a cross-sectional dataset for the year 2005. For this, we have to assume a specific general equilibrium structure, and we rely on the one proposed by Anderson and van Wincoop (2003) for convenience.

The obtained results suggest that ignoring endogenous selection into PTAs is relatively harmful. The impact of endogenous PTAs on members' relative to nonmembers' trade flows is about 188 percentage points higher than in a model which assumes PTA membership to be exogenous. With the dataset at hand, the process of zero versus positive exports should be modeled separately from the one of positive exports. Ignoring the latter leads to a downward bias of the predicted trade effects of PTAs by about 73 percentage points as compared to the preferred model.

Appendix: Country coverage (126 economies)

The following set of countries is covered in our dataset:

Albania, Algeria, Argentina, Armenia, Australia, Austria, Azerbaijan, Bangladesh, Barbados, Belarus, Belgium, Belize, Benin, Bolivia, Bosnia and Herzegovina, Botswana, Brazil, Brunei, Bulgaria, Burundi, Cameroon, Canada, Chile, China, Colombia, Comoros, Costa Rica, Cote d'Ivoire, Croatia, Cyprus, Czech Republic, Denmark, Ecuador, Egypt, El Salvador, Estonia, Ethiopia, Fiji, Finland, France, Georgia, Germany, Ghana, Greece, Guatemala, Guinea, Honduras, Hungary, Iceland, India, Indonesia, Islamic Rep. Iran, Ireland, Israel, Italy, Jamaica, Japan, Jordan, Kazakhstan, Kenya, Rep. Korea, Kuwait, Kyrgyz Republic, Latvia, Lebanon, Lithuania, Luxembourg, FYR Macedonia, Madagascar, Malawi, Malaysia, Malta, Mauritius, Mexico, Moldova, Morocco, Mozambique, Namibia, Nepal, Netherlands, New Zealand, Nicaragua, Niger, Nigeria, Norway, Oman, Pakistan, Panama, Papua New Guinea, Paraguay, Peru, Philippines, Poland, Portugal, Qatar, Romania, Russian Federation, Saudi Arabia, Senegal, Singapore, Slovak Republic, Slovenia, South Africa, Spain, Sri Lanka, St. Lucia, Sudan, Suriname, Sweden, Switzerland, Syrian Arab Republic, Tajikistan, Tanzania, Thailand, Togo, Trinidad and Tobago, Tunisia, Turkey, Turkmenistan, Uganda, United Kingdom, United States, Uruguay, Vanuatu, RB Venezuela, Zimbabwe.

PTA coverage (121 agreements).—Our dataset includes all PTAs notified to the World Trade Organization that are active since 2005 or earlier. The data are augmented and corrected by using information from PTA secretariat web-pages. This leads to a coverage of the following PTAs in our dataset:

ASEAN Free Trade Area (AFTA), Albania and Bosnia and Herzegovina, Albania and Bulgaria, Albania and FYR Macedonia, Albania and Moldova, Albania and Romania, Armenia and Kazakhstan, Armenia and Moldova, Armenia and Russian Federation, Armenia and Turkmenistan, Association of Southeast Asian Nations (ASEAN), Baltic Free Trade Area (BAFTA), Bangkok Agreement, Bulgaria and Bosnia and Herzegovina, Bulgaria and FYR Macedonia, Bulgaria and Israel, Bulgaria and Turkey, Central American Common Market (CACM), Andean Subregional Integration Agreement (Cartagena Agreement, CAN), Canada and Chile, Canada and Israel, Canada and Costa Rica, Caribbean Community (CARICOM), Central European Free Trade Agreement (CEFTA), Australia New Zealand Closer Economic Relations Trade Agreement (CER), Chile and Costa Rica, Chile and El Salvador, Chile and Mexico, Commonwealth of Independent States Free Trade Agreement (CIS), Common Market for Eastern and Southern Africa (COMESA), Croatia and Albania, Croatia and Bosnia and Herzegovina, Croatia and FYR Macedonia, East African Community Treaty (EAC), Eurasian Economic Community (EAEC), European Community (EC), EC and Algeria, EC and Bulgaria, EC and Chile, EC and Croatia, EC and Egypt, EC and FYR Macedonia, EC and Iceland, EC and Israel, EC and Jordan, EC and Lebanon, EC and Mexico, EC and Morocco, EC and Norway, Economic Cooperation Organization (ECO), EC and Romania, EC and South Africa, EC and Switzerland and Liechtenstein, EC and Syria, EC and Tunesia, EC and Turkey, Agreement on the European Economic Area (EEA), European Free Trade Association (EFTA), EFTA and Bulgaria, EFTA and Chile, EFTA and Croatia, EFTA and FYR Macedonia, EFTA and Israel, EFTA and Jordan, EFTA and Mexico, EFTA and Morocco, EFTA and Romania, EFTA and Singapor, EFTA and Tunisia, EFTA and Turkey, FYR Macedonia and Bosnia and Herzegovina, The Unified Economic Agreement between the Countries of the Gulf Cooperation Council (GCC), Georgia and Armenia, Georgia and Kazakhstan, Georgia and Russian Federation, Georgia and Turkmenistan, Global System of Trade Preferences among Developing Countries (GSTP), India and Sri Lanka, Israel and Turkey, Japan and Mexico, Japan and Singapor, Kyrgyz Republic and Armenia, Kyrgyz Republic and Kazakhstan, Kyrgyz Republic and Moldova, Kyrgyz Republic and Russian Federation, Asociación Latinoamericana de Integración (ALADI, LAIA), Mercado Común del Sur (MERCOSUR), Mexico and Israel, Moldova and Bosnia and Herzegovina, Moldova and Bulgaria, Moldova and Croatia, Moldova and FYR Macedonia, Melanesian Spearhead Group Free Trade Area Agreement (MSG), North American Free Trade Agreement (NAFTA), New Zealand and Singapore, Panama and El Salvador, Papua New Guinea–Australia Trade and Commercial Relations Agreement (PATCRA), Protocol relating to Trade Negotiations among Developing Countries (PTN), Rep. of Korea and Chile, Romania and Bosnia and Herzegovina, Romania and FYR Macedonia, Romania and Israel, Romania and Moldova, Romania and Turkey, Southern African Development Community (SADC), South Asian Association for Regional Cooperation Preferential Trading Arrangement (SAPTA), Singapore and Australia, South Pacific Regional Trade and Economic Cooperation Agreement (SPARTECA), Thailand and Australia, TRIPARTITE, Turkey and Bosnia and Herzegovina, Turkey and Croatia, Turkey and FYR Macedonia, United States and Chile, United States and Isreal, United States and Jordan, United States and Singapore, Unites States and Australia, Traite Modifié de l'Union Économique et Monétaire Ouest Africaine (WAEMU/UEMOA).

REFERENCES

Abadie, Alberto, and Guido W. Imbens. 2006. "Large Sample Properties of Matching Estimators for Average Treatment Effects." *Econometrica*, 74(1): 235–67.

Aitken, Norman D. 1973. "The Effect of the EEC and EFTA on European Trade: A Temporal Cross-Section Analysis." *American Economic Review*, 63(5): 881–92.

Anderson, James E. 1979. "A Theoretical Foundation for the Gravity Equation." *American Economic Review*, 69(1): 106–16.

Anderson, James E. 2009. "Gravity, Productivity and the Pattern of Production and Trade." National Bureau of Economic Research Working Paper 14642.

Anderson, James E., and Eric van Wincoop. 2002. "Borders, Trade, and Welfare." In *Brookings Trade Forum 2001*, ed. Susan M. Collins and Dani Rodrik, 207–43. Washington, DC: Brookings Institution.

Anderson, James E., and Eric van Wincoop. 2003. "Gravity with Gravitas: A Solution to the Border Puzzle." *American Economic Review*, 93(1): 170–92.

Baier, Scott L., and Jeffrey H. Bergstrand. 2002. "On the Endogeneity of International Trade Flows and Free Trade Agreements." Unpublished.

Baier, Scott L., and Jeffrey H. Bergstrand. 2004. "Economic Determinants of Free Trade Agreements." *Journal of International Economics*, 64(1): 29–63.

Baier, Scott L., and Jeffrey H. Bergstrand. 2007. "Do Free Trade Agreements Actually Increase Members' International Trade?" *Journal of International Economics*, 71(1): 72–95.

Baier, Scott L., and Jeffrey H. Bergstrand. 2009. "Estimating the Effects of Free Trade Agreements on International Trade Flows Using Matching Econometrics." *Journal of International Economics*, 77(1): 63–76.

Baldwin, Richard. 2009. "Big-Think Regionalism: A Critical Survey." In *Regional Rules in the Global Trading System*, ed. Antoni Estevadeordal, Kati Suominen, and Robert Teh, 17–95. Cambridge, UK: Cambridge University Press.

Baldwin, Richard E., and Anthony J. Venables. 1995. "Regional Economic Integration." In *Handbook of International Economics*. Vol. 3, ed. Gene M. Grossman and Kenneth Rogoff, 1597–1644. Amsterdam: Elsevier.

Bitler, Marianne P., Jonah B. Gelbach, and Hilary W. Hoynes. 2006. "What Mean Impacts Miss: Distributional Effects of Welfare Reform Experiments." *American Economic Review*, 96(4): 988–1012.

Cameron, Colin A., and Pravin K. Trivedi. 2005. *Microeconometrics, Methods and Applications*. New York: Cambridge University Press.

Carrère, Céline. 2006. "Revisiting the Effects of Regional Trade Agreements on Trade Flows with Proper Specification of the Gravity Model." *European Economic Review*, 50(2): 223–47.

Cragg, John G. 1971. "Some Statistical Models for Limited Dependent Variables with Application to the Demand for Durable Goods." *Econometrica*, 39(5): 829–44.

Dixit, Avinash K., and Joseph E. Stiglitz. 1977. "Monopolistic Competition and Optimum Product Diversity." *American Economic Review*, 67(3): 297–308.

Dow, William H., and Edward C. Norton. 2003. "Choosing Between and Interpreting the Heckit and Two-Part Models for Corner Solutions." *Health Services and Outcomes Research Methodology*, 4(1): 5–18.

Duan, Naihua, Willard G. Manning, Jr., Carl N. Morris, and Joseph P. Newhouse. 1984. "Choosing between the Sample-Selection Model and the Multi-part Model." *Journal of Business and Economic Statistics*, 2(3): 283–89.

Eaton, Jonathan, and Samuel Kortum. 2002. "Technology, Geography, and Trade." *Econometrica*, 70(5): 1741–79.

Eaton, Jonathan, and Akiko Tamura. 1994. "Bilateralism and Regionalism in Japanese and U.S. Trade and Direct Foreign Investment Patterns." *Journal of the Japanese and International Economies*, 8(4): 478–510.

Egger, Hartmut, Peter Egger, and David Greenaway. 2008. "The Trade Structure Effects of Endogenous Regional Trade Agreements." *Journal of International Economics*, 74(2): 278–98.

Egger, Peter, Mario Larch, Kevin E. Staub, and Rainer Winkelmann. 2011. "The Trade Effects of Endogenous Preferential Trade Agreements: Dataset." *American Economic Journal: Economic Policy*, http://www.aeaweb.org/articles.php?doi=10.1257/pol.3.3.113.

Feenstra, Robert C. 2004. *Advanced International Trade: Theory and Evidence*. Princeton: Princeton University Press.

Feenstra, Robert, and Hiau Looi Kee. 2008. "Export Variety and Country Productivity: Estimating the Monopolistic Competition Model with Endogenous Productivity." *Journal of International Economics*, 74(2): 500–518.

Felbermayr, Gabriel J., and Wilhelm Kohler. 2006. "Exploring the Intensive and Extensive Margins of World Trade." *Review of World Economics*, 142(4): 642–74.

Freund, Caroline. 2000. "Different Paths to Free Trade: The Gains from Regionalism." *Quarterly Journal of Economics*, 115(4): 1317–41.

Glejser, Herbert. 1968. "An Explanation of Differences in Trade-Product Ratios among Countries." *Cahiers Economiques de Bruxelles*, 37: 47–58.

Greene, William H. 2008. *Econometric Analysis*. 6th ed. Upper Saddle River, NJ: Prentice Hall.

Greenaway, David, and Chris Milner. 2002. "Regionalism and Gravity." *Scottish Journal of Political Economy*, 49(5): 574–85.

Heckman, James J., Lance Lochner, and Christopher Taber. 1998. "General-Equilibrium Treatment Effects: A Study of Tuition Policy." *American Economic Review*, 88(2): 381–86.

Helpman, Elhanan, Marc Melitz, and Yona Rubinstein. 2008. "Estimating Trade Flows: Trading Partners and Trading Volumes." *Quarterly Journal of Economics*, 123(2): 441–87.

Kehoe, Timothy J., and Kim J. Ruhl. 2009. "How Important is the New Goods Margin in International Trade?" Federal Reserve Bank of Minneapolis Staff Report 324.

Krugman, Paul. 1980. "Scale Economies, Product Differentiation, and the Pattern of Trade." *American Economic Review*, 70(5): 950–59.

Leung, Siu Fai, and Shihti Yu. 1996. "On the Choice between Sample Selection and Two-Part Models." *Journal of Econometrics*, 72(1–2): 197–229.

Magee, Christopher S. 2003. "Endogenous Preferential Trade Agreements: An Empirical Analysis." *Contributions to Economic Analysis and Policy*, 2(1).

Manning, Willard G., Naihua Duan, and W. H. Rogers. 1987. "Monte Carlo Evidence on the Choice between Sample Selection and Two-Part Models." *Journal of Econometrics*, 35(1): 59–82.

Marshall, Monty G., and Keith Jaggers. 2007. *Polity IV Project: Political Regime Characteristics and Transitions, 1800–2006*. College Park, MD: Center for Systemic Peace.

McDonald, John F., and Robert A. Moffitt. 1980. "The Uses of Tobit Analysis." *Review of Economics and Statistics*, 62(2): 318–21.

Melitz, Marc J., and Giancarlo I. P. Ottaviano. 2008. "Market Size, Trade, and Productivity." *Review of Economic Studies*, 75(1): 295–316.

Monfardini, Chiara, and Rosalba Radice. 2008. "Testing Exogeneity in the Bivariate Probit Model: A Monte Carlo Study." *Oxford Bulletin of Economics and Statistics*, 70(2): 271–82.

Rodrik, Dani. 1995. "Political Economy of Trade Policy." In *Handbook of International Economics*. Vol. 3, ed. Gene M. Grossman and Kenneth Rogoff, 1457–94. Amsterdam: Elsevier.

Santos Silva, J. M. C., and Silvana Tenreyro. 2006. "The Log of Gravity." *Review of Economics and Statistics*, 88(4): 641–58.

Santos Silva, J. M.C., and Silvana Tenreyro. 2008. "Trading Partners and Trading Volumes: Implementing the Helpman-Melitz-Rubinstein Model Empirically." University of Essex, Department of Economics Discussion Paper 662.

Soloaga, Isidro, and L. Alan Winters. 2001. "Regionalism in the Nineties: What Effect on Trade?" *North American Journal of Economics and Finance*, 12(1): 1–29.

Terza, Joseph V. 1998. "Estimating Count Data Models with Endogenous Switching: Sample Selection and Endogenous Treatment Effects." *Journal of Econometrics*, 84(1): 129–54.

Tinbergen, Jan. 1962. *Shaping the World Economy*. New York: Twentieth Century Fund.

Windmeijer, F. A. G., and J. M. C. Santos Silva. 1997. "Endogeneity in Count Data Models: An Application to Demand for Health Care." *Journal of Applied Econometrics*, 12(3): 281–94.

Wooldridge, Jeffrey M. 2002. *Econometric Analysis of Cross Section and Panel Data*. Cambridge, MA: MIT Press.

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