The Trade Effects of Endogenous Preferential Trade Agreements*

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Abstract

This paper suggests an empirical modeling strategy which takes full account of the structural, non-(log-)linear impact of endogenous trade barriers on trade in new trade theory models. In particular, we use the framework to evaluate and quantify the role of endogenous preferential trade agreement (PTA) membership for trade. Apart from paying attention to structural modeling of the impact of trade policy on trade, the suggested model takes into account both that preferential trade agreement 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.

Key words: Gravity model; Endogenous preferential trade agreement membership; Poisson pseudo-maximum likelihood estimation with endogenous binary indicator variables

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1 Introduction

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 Baier and Bergstrand, 2002, 2004, 2009; Magee, 2003; Egger, Egger, and Greenaway, 2008): PTAs are most likely concluded among large, similarly-sized, non-distant economies which are relatively autocratic and 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.

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 Wooldridge, 2002; Cameron and Trivedi, 2005). Obviously, this is at odds with general equilibrium. Heckman, Lochner, and 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. 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

[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 Rodrik (1995), Baldwin and Venables (1995), and Baldwin (2008), for details.

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 Eaton and Kortum, 2002; Anderson and 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 Santos Silva and Tenreyro, 2006; 2008; and Helpman, Melitz, and 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 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 data-set 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 reduces trade flows among members directly, but it entails also 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 potential

[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 Tinbergen (1962), Glejser (1968), Aitken (1973), for some of the earliest examples and Freund (2000), Soloaga and Winters (2001), and Carrère (2006) for more recent ones. Greenaway and 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.

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. For instance, a log-linear model of exports which ignores general equilibrium effects on top of the other problems leads to a bias of -73 percentage points or -66% relative to the preferable two-part PPML approach. A one-part Poisson pseudo-maximum likelihood (PPML) 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 -56 percentage points or -51% relative to a two-part PPML 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 -11 percentage points. As compared to these biases it is less harmful to ignore that PTA membership effects are heterogeneous due to the variation in most-favored nation tariff rates. For instance, ignoring heterogeneous tariffs in the preferable two-part PTA model leads to a downward bias of the PTA-induced effect of less than one-fifth 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 3 points out three problems with the implementation of that model in applied work targeted towards the analysis of PTA membership effects on trade. Section 4 describes the specification and data. Section 5 introduces the modeling strategy to overcome these obstacles by treating zero trade flows implicitly, and presents the corresponding estimation results. Section 6 derives a zero-inflated gravity equation, lays out the econometric two-part model, and gives the estimation results thereof. Section 7 computes the impact of PTA membership as observed in the year 2005 to a situation without any PTA memberships in the same year. The last section concludes with a summary of the most important findings.

2 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 Krugman (1980) with love-of-variety preferences à la Dixit and Stiglitz (1977). Their framework can be briefly introduced as follows. Let us denote nominal exports of country i to country j (with $i, j = 1, ..., N$) 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 solutions of the following set of $2N$ equations[3]

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

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*

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

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 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

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

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

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

where $Z_{ij} = (1, \ln DIST_{ij}, BORD_{ij}, ...)$ is a vector containing a constant and all trade cost or trade facilitating variables except PTA_{ij}. Generally, binary variables such as $BORD_{ij}$ enter as they are in Z_{ij} and continuous variables such as $DIST_{ij}$ enter in logs as in (3). Moreover, $\beta = (\beta_0, \beta_1, \beta_2, \ldots)$ is a vector of coefficients corresponding to the elements in 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 $\beta_0 = -\ln y_W$. Moreover, the multilateral resistance terms Π_i and P_j are determined as in (2), and thus implicit functions of t_{ij}.

3 Empirical problems with the implementation of a structural gravity model

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

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

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). For estimation of the parameters β and δ in the empirical model (5), α_i and γ_j may be captured by fixed country effects. Given these parameters, the $2N$ multilateral resistance terms in (2) may be computed subsequently. Estimation of β and δ does not hinge upon the general equilibrium structure of the model,[4] and it is well-known that the estimation part of the problem covers a wide range of (one-sector) models such as the multi-country version of the Dixit-Stiglitz-Krugman model, Eaton and Kortum (2002), or 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 will influence 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).[5]

[4] Notice that general equilibrium effects are fully captured by the country fixed effects in estimation.

[5] Previous work predominantly relied on Heckman-type switching regression models (Baier and

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 structural models such as the one of Anderson and van Wincoop (2003). We will show how model (1) can be adapted to account for some endogenous trade frictions, still obeying (2). Obviously, such a goal can only be achieved by means of instrumental variable estimation.

Second, depending on the data-set in use, the $N(N-1)$-size vector $\mathbf{X}$ 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 $\mathbf{X}$, 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, making multiplicative estimation of the model even more attractive.

We elaborate on these issues in Section 5, 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.

4 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:

Bergstrand, 2002; Magee, 2003) or matching methods based on the propensity score (Baier and Bergstrand, 2002, 2009).

variables capturing political affinities or impediments to bilateral trade liberalization; country size and relative factor endowments; and proxies for iceberg trade costs. We classify two countries as belonging to a common PTA, if they are active since 2005 or earlier as notified to the World Trade Organization. The data are augmented and corrected by using information from PTA secretariat web-pages and they are 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 and it is reasonable to assume that they do not affect trade flows directly. The associated variables are based on the data collected in the Polity IV Project (see Marshall and Jaggers, 2007). In particular, we include the absolute difference in a score variable, measuring the autocracy of an exporter and an importer, respectively ($AUTOC_{ij}$);[6] the squared value of the latter variable ($AUTOC_{ij}^2$); the absolute difference in a variable, measuring the durability of an exporter's and an importer's political regime, respectively ($DURAB_{ij}$);[7] the squared value of the latter variable ($DURAB_{ij}^2$); the absolute difference in a score variable, measuring the political competition in the government of an exporter and an importer, respectively ($POLCOMP_{ij}$);[8] the squared value of the latter variable ($POLCOMP_{ij}^2$).

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 de-

[6] $AUTOC$ 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.

[7] $DURAB$ 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. $DURAB$ is computed for all years beginning with the first regime change since 1800 or the date of independence if that event occurred after 1800.

[8] $POLCOMP$ 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.

terminants 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. Accordingly, we include a variable measuring the total bilateral real GDP, $\text{RGDPsum}_{ij} = log(\text{RGDP}_i + \text{RGDP}_j)$ with RGDP_i and RGDP_j denoting the real GDP of country i and j, respectively. Similarity of two countries' size in terms of GDP is defined as $\text{RGDPsim}_{ij} = log\{1 - [\text{RGDP}_i/(\text{RGDP}_i + \text{RGDP}_j)]^2 - [\text{RGDP}_j/(\text{RGDP}_i + \text{RGDP}_j)]^2\}$. The probability of a bilateral PTA membership between countries i and j is expected to rise with RGDPsim_{ij}.

Moreover, Baier and Bergstrand (2004) 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 (DKL_{ij}) and the other one captures the capital-labor relative factor endowment difference between that pair and the rest of the world (DROWKL_{ij}). In our application, the two variables are defined as follows: $\text{DKL}_{ij} = |log(\text{RGDP}_i/\text{POP}_i) - log(\text{RGDP}_j/\text{POP}_j)|$, where $\text{RGDP}_i/\text{POP}_i$ measures country i's real GDP per capita; $\text{DROWKL}_{ij} = 0.5\{|log(\sum_{k\neq i}\text{RGDP}_k/\sum_{k\neq i}\text{POP}_k) - log(\text{RGDP}_i/\text{POP}_i)| + |log(\sum_{k\neq j}\text{RGDP}_k/\sum_{k\neq j}\text{POP}_k) - log(\text{RGDP}_j/\text{POP}_j)|\}$.[9] Following Baier and Bergstrand, we expect the probability of bilateral PTA membership to rise with DKL_{ij} and to fall with DROWKL_{ij}. Data on real GDP and population are taken from the World Bank's World Development Indicators.

Proxies for iceberg trade costs: Log bilateral (great circle) distance between two countries' capitals (DIST_{ij});[10] the squared log distance to capture a higher degree of non-linearity in geographical distance space (DIST^2_{ij});[11] an indicator variable which is set to one if

[9]Notice that Baier and Bergstrand employ capital-labor ratios while we have to use real GDP per capita instead for reasons of data availability (the data-set used here contains 15,750 country-pairs while the one in Baier and Bergstrand (2004) covered only 1,453 country-pairs). However, capital-labor ratios are highly correlated with real GDP per capita.

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

[11]Notice that the inclusion of DIST^2_{ij} substitutes for an indicator variable which is one in case of a common land border between countries i and j and zero else in the application. Including DIST^2_{ij} and such an indicator together renders the parameter of the latter insignificant.

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 one of 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'Etudes Prospectives et d'Informations Internationales (CEPII). The list of variables in Baier and Bergstrand (2004) did not include (country dummies and) DIST^2_{ij}, LANG_{ij}, COLONY_{ij}, CURCOL_{ij}, COMCOL_{ij}, or SMCTRY_{ij}. We only include a subset of these variables in the exports outcome equation since the other ones do not display a significant direct impact on exports.[12]

In some of the econometric models applied here, selection into positive exports has a stochastic component and is otherwise determined by a function of a complete set of exporter and importer dummy variables and the following set of regressors: the PTA indicator variable; log bilateral distance between two countries' capitals (DIST_{ij}); the aforementioned common language indicator (LANG_{ij}); and an indicator variable which is set to one if two countries have a common land border and zero else (BORD_{ij}).[13] Whenever both selection into positive exports and into PTAs are specified in the mentioned way, we model the two processes as a recursive bivariate probit model.

Finally, in our application we include the following trade cost variables in Z_{ij} in the nominal exports outcome equation (5): DIST_{ij}, BORD_{ij}, and LANG_{ij}. Otherwise, nomi-

[12]See Helpman, Melitz, and Rubinstein (2008) for a similar approach.

[13]As mentioned before, the impact of a border indicator variable may be thought of as a non-log-linear impact of distance in the right-hand-side specification of the selection model. We employ it here instead of the squared distance variable DIST^2_{ij}, since this specification works better than one that exhibits a right-hand side of the zero-versus-positive exports hurdle model which is more similar to the right-hand side of the selection-into-PTAs model.

nal exports are a function of a complete set of exporter and importer dummy variables,[14] and of (potentially endogenous) PTA_{ij}. Data on bilateral exports in nominal U.S. dollars are collected from the United Nation's World Trade Database.

−− Table 1 −−

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

5 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 2. 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 this end is exemplified in Section 7, our objective in the subsequent sections is to consistently estimate β and δ.

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

5.1 Econometric model

Since the parameters of interest in model (5) are β and δ, terms α_i and γ_j can be considered as nuisance parameters from an econometric point of view. 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 to be used depend on the assumptions on the relationship between (α_i, γ_j) and the regressors, Z_{ij} and PTA_{ij}. If (α_i, γ_j) were independent of Z_{ij} and PTA_{ij}, random effects estimation would be consistent and efficient. However, as independence is precluded by the underlying economic model which suggests that α_i and γ_j depend on Z_{ij} and PTA_{ij}, the model should be treated as a two-way fixed effects model and is equivalent to a model with a comprehensive set of exporter and importer dummies.

There are two important differences to a standard panel data model, though. First, this model is non-linear, making it impossible to use simple transformations to eliminate the fixed effects. Second, since the data consist of all possible pairs of N countries, and countries take on both roles, exporters and importers, 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 no special adjustment to the estimation methods is required.[15] Accordingly, the country-specific components can be estimated analogously to the linear fixed effects model by including a dummy variable for each importer and exporter country. This procedure is computationally intensive, given 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) to be estimated is

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

[15] 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.

Under the assumption of exogenous PTA membership, $\mathrm{E}(\epsilon_{ij}|Z_{ij}, PTA_{ij}, \alpha_i, \gamma_j) = 1$ 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},

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

where 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 W_{ij} have been listed in Section 4 and they contain elements of 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 Terza (1998), it is possible to derive a tractable form of $\mathrm{E}[X_{ij}|Z_{ij}, PTA_{ij}, W_{ij}, \alpha_i, \gamma_j]$ under the assumption of bivariate normality of v_{ij} and $\ln(\epsilon_{ij})$, which leads to the following expressions

$$\mathrm{E}[X_{ij}|Z_{ij}, PTA_{ij}, W_{ij}, \alpha_i, \gamma_j] = \lambda_{ij}\Psi_{ij}, \tag{8}$$

with

$$\begin{aligned} \lambda_{ij} &\equiv \exp[Z'_{ij}\beta + \delta PTA_{ij} + \alpha_i + \gamma_j] \quad \text{and} \qquad (9) \\ \Psi_{ij} &\equiv \mathrm{E}[\epsilon_{ij}|Z_{ij}, PTA_{ij}, W_{ij}, \alpha_i, \gamma_j] \\ &= PTA_{ij}\frac{\Phi(\vartheta + W'_{ij}\theta)}{\Phi(W'_{ij}\theta)} + (1 - PTA_{ij})\frac{1 - \Phi(\vartheta + W'_{ij}\theta)}{1 - \Phi(W'_{ij}\theta)}. \end{aligned}$$

The last equality follows from joint normality of the errors, where $\Phi(\cdot)$ denotes the cumulative distribution function of the standard normal distribution.[16] The parameter ϑ is

[16] An alternative estimation technique which does not rely on bivariate normality is the GMM approach of 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 6.2.

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 $\mathrm{E}(\epsilon_{ij}|Z_{ij}, PTA_{ij}, \alpha_i, \gamma_j) = 1$. However, if $\rho \neq 0$, estimation of the parameters β contained in λ_{ij} will be inconsistent if Ψ_{ij} is neglected.

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 $\mathrm{E}(\epsilon_{ij}|Z_{ij}, PTA_{ij}, \alpha_i, \gamma_j) = 1$ (Santos Silva and Tenreyro, 2006). 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.

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,[17] the results reported in Santos Silva and Tenreyro (2006) strongly encourage us towards viewing Poisson PML estimates as more efficient.

As a practical matter, we estimate (8) in two steps, as this is easy to do from a computational angle. First, estimation of (7) is carried out by Probit regression, which yields

[17] 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 $Z'_{ij}\beta + \delta PTA_{ij} + \alpha_i + \gamma_j$, 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.

estimates $\hat{\theta}$. Using these in (8) for θ, we optimize over β, δ and ϑ. As a consequence of applying two-step procedures, second-step standard errors have to be adjusted to account for the variance of first-step estimates.[18]

5.2 Estimation results

It is the aim of this section to apply the aforementioned methods to estimate the parameters needed to infer the impact of endogenous PTA membership on exports while allowing for zero exports in the data-generating process. In this subsection, we summarize the parameter estimates from PPML and NLS models described in Section 5.1. Table 2 displays the parameter estimates of five alternative models of nominal bilateral exports in U.S. dollars (X_{ij}). In the second column, we take log exports as the dependent variable and report the parameters of the four covariates of interest in the export equation – PTA_{ij}, DIST_{ij}, BORD_{ij}, and LANG_{ij} – estimating a log-linear model via OLS and treating PTA_{ij} as exogenous. In columns three and four, we report parameters with both PPML and NLS when treating PTA_{ij} as exogenous. In columns five and six, we treat PTA_{ij} as endogenous for both PPML and NLS. In the latter case, we use a first stage probit model based on the covariates mentioned in Section 4 which obtains parameters that are summarized in Table 3. This assumes that political variables, bilateral size, bilateral endowments, and measures of iceberg trade costs, which serve as identifying instruments, are exogenous.[19] Hence, countries select into PTA membership under favorable political and economic circumstances which – after controlling for other determinants of trade flows – do not directly affect trade.

–– Tables 2 and 3 ––

[18]Details for the NLS variance estimator are given in Terza (1998). As the form of the Poisson PML variance estimator is very similar to its NLS variant, we dispense with its exposition.

[19]Notice that size, factor endowment, and trade cost variables may only be used as identifying instruments to the extent that they explain PTA membership beyond fixed country effects and exogenous trade cost variables in the trade flow model. Fixed country effects capture the influence of unilateral determinants of GDP and prices comprehensively in a model based on the system in (1) and (2).

The results in Table 2 suggest the following conclusions. First of all, as the discussion below indicates, selection into PTAs based on observables is positive. Observed factors raising the probability of joining a PTA also have a trade-increasing effect. Hence, particularly those country-pairs which display a high level of goods trade flows anyway select into PTAs. 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.

Second, there is evidence for selection 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 in the PPML model, thus rejecting exogeneity of PTA_{ij}.[20] A negative ϑ indicates that unobservables (i.e., 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 parameters: The point estimate for PTA_{ij} increases as we abandon the assumption of PTA_{ij} to be exogenous. This is true for OLS, PPML and NLS. Not surprisingly, the major difference across columns for PPML and NLS estimates, respectively, arises for the parameter of PTA_{ij}. The remaining parameters are fairly similar across the columns. However, the estimates differ relatively starkly between PPML and NLS. Yet, there we know that PPML is preferable over NLS according to the discussion in Santos Silva and Tenreyro (2006) and Section 5.1 above.

The results from the probit estimation for the reduced form equation of PTA suggest the following conclusions. The political variables turn out to be important for the decision to form or join a PTA as in Egger, Egger, and Greenaway (2008). Specifically, the durability of an exporter's and an importer's political regime turns out to influence the

[20]The point estimate for $\hat{\vartheta}$ based on NLS is similar to PPML's, although it is only borderline significant. Since NLS is less efficient (Santos Silva and Tenreyro, 2006), an endogeneity test based on the PPML estimate has more power and should be preferred over the NLS's.

probability to conclude a PTA positive at the mean of 29.4 ($0.0059 \times \text{DURAB}_{ij} - 0.0001 \times \text{DURAB}^2_{ij} = 0.1705$). The political competition index ($POLCOMP_{ij}$) as well as the autocracy index ($AUTOC_{ij}$) turn out to exert a non-linear effect on the probability to form a PTA. Whereas an increase in political competition reduces the latent variable determining PTA membership at low values of the political competition index, high values imply a higher value of the latent variable behind PTA membership. On the contrary, a marginal increase of the autocracy index exerts a positive influence on the latent variable underlying PTA membership at low values of the index and a negative influence at high values of the index. Neither DURAB_{ij} nor $POLCOMP_{ij}$ or $AUTOC_{ij}$ were included in the models of Magee (2003) or Baier and Bergstrand (2004).

Distance has a negative effect on the probability to conclude a PTA. Even though the coefficient of DIST_{ij} is positive, the marginal effect is negative for all observations, since for the minimum value of DIST_{ij} of 3.25, the overall impact on the latent variable associated with PTA membership is equal to $0.2332 \times \text{DIST}_{ij} - 0.0812 \times \text{DIST}^2_{ij} = -0.0998$. This result is consistent with the results of Magee (2003) – who found a negative impact of log distance on PTA membership in a cross-section of 4,786 country-pairs and a similar effect with panel data – and of Baier and Bergstrand (2004) – who found a negative impact of log distance in a cross-section of 1,431 pairings.

The capital-labor relative factor endowment difference between two countries i and j exerts a negative impact on the probability of PTA membership of i and j. In Magee's (2003) cross-sectional models, the impact of this variable does not affect PTA membership significantly. Baier and Bergstrand (2004) found that capital-labor ratio differences affected PTA membership significantly positively. The capital-labor relative factor endowment difference between pair ij and the rest of the world affects the probability of i and j to be members of the same PTA positively, unlike in Baier and Bergstrand (2004).[21]

[21] Note that Baier and Bergstrand (2004) were able to use a better measure of capital-labor ratios in their much smaller sample of countries than we are able to do here. However, a comparison of our results with theirs and those of Magee (2003) is difficult, since they did not include fixed exporter and importer effects (and some other control variables that we employ) in their cross-sectional models.

Among the effects of cultural, geographical, and political indicator variables, the ones of common language LANG_{ij} and COLONY_{ij} are statistically insignificant. The effect of LANG_{ij} on PTA membership in 1998 in Magee's (2003) application was positive and significant. Both COLONY_{ij} and LANG_{ij} were absent from Baier and Bergstrand's (2004) models. However, we find statistically significant effects of a positive influence if countries are on the same continent CONT_{ij} (consistent with Baier and Bergstrand, 2004), if they had a common colonizer, COMCOL_{ij}, if one of them was a colony of the other after 1945, CURCOL_{ij}, and if one country was part of the other in the past, SMCTRY_{ij}. These variables were not included in the specifications of Magee (2003) and Baier and Bergstrand (2004). Interestingly, these variables do not matter significantly on their own when conditioning on the control variables in the trade flow equation and the fixed country effects.

6 Modeling zero trade flows explicitly

The previous approach accommodated 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 Felbermayr and Kohler, 2006) to allow for log-linearization. However, the aforementioned models assumed that zero exports were proportionally generated by the stochastic processes at stake. It is not advisable to use the methods discussed before with a large mass of zeros in the data. With bilateral trade matrices, the problem of large numbers of zeros is well documented (see Felbermayr and Kohler, 2006; and Helpman, Melitz, and Rubinstein, 2008). Beyond econometric issues, it may be interesting to distinguish between the effect of PTA membership on the extensive country margin of exports – i.e., the number of pairings which started exporting because of PTA membership – relative to the intensive margin – the extent to which PTA membership raised exports among pairs that traded already.

Before turning to the econometric modeling of zero-inflated gravity equations, let us return to the theoretical model introduced in Section 2 and augment it so as to allow for zero trade flows in the deterministic part of the model. We will do so by introducing decisions of symmetric monopolistically competitive firms as in Krugman (1980) in each country, where the extent of fixed bilateral market entry costs relative to operating profits in that market governs a firm's decision to serve the target market via exports or not.[22]

6.1 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 $j = 1, ..., N$ 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 that all producers in country i are symmetric with respect to $\hat{z}_j(b)$ and $f_j(b)$. 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 $x_{ij} \equiv c_{ij} t_{ij} = p_i^{-\sigma} t_{ij}^{1-\sigma} P_j^{\sigma-1} y_j$.

[22] This reasoning is not novel. For instance, the source of zero trade flows in Helpman, Melitz, and Rubinstein (2008) is the same as it will be below. Unlike they do, we will not venture into modeling firms as heterogeneous in terms of productivity for the sake of brevity and to focus on the empirical issue at stake. However, it is nevertheless useful to outline the model to make transparent how the econometric model needs to be changed and what can be learned for the impact of PTAs on bilateral exports. Also, for an illustration of the comparative static effects of preferential trade liberalization we need to specify a general equilibrium structure, even though it could be different from the one applied here.

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 $\mathbb{I}_{ij}$ 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

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

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

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

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:

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

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

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

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.

6.2 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):

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

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

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

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 (Cragg, 1971; Duan, Manning, Morris, and Newhouse, 1983) have been discussed in econometrics for some time, 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,

$$\mathrm{E}(X_{ij}|Z_{ij}, W_{ij}, PTA_{ij}, \mathbb{I}_{ij} = 1) = \lambda_{ij}\Psi_{ij}, \tag{18}$$

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 5.

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 $\mathbb{I}_{ij}$ as defined by equation (11) is translated into a stochastic process

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

where the vector 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 Q_{ij} may but need not contain the same elements as 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 ($\mathbb{I}_{ij}$) 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. Monfardini and Radice, 2008, for some recent applications). Then, the probability of trading conditional on PTA membership can be written as (see, e.g., Greene, 2008)

$$\begin{aligned} &\Pr(\mathbb{I}_{ij} = 1|Q_{ij}, W_{ij}, PTA_{ij}) = \\ &\frac{\Phi_2[(2PTA_{ij} - 1)W'_{ij}\theta, Q'_{ij}\omega + \kappa PTA_{ij}, (2PTA_{ij} - 1)\rho_{v\xi}]}{\Phi[(2PTA_{ij} - 1)W'_{ij}\theta]}, \end{aligned} \tag{20}$$

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 – but to have small impact on the extensive margin – an increase of the value of positive bilateral exports resulting principally from β.

A convenience of such a model is that the two parts, (18) and (20), can be estimated independently. 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 to the estimation of parameters in (18), we can use the same two-stage PML or NLS procedures described in Section 5, and include only the observations with positive exports in the estimation.

The two-part model presented in this section differs from the Heckman-type sample-selection models suggested in the recent literature for discriminating between effects at the extensive and intensive margins of trade with exogenous regressors (see Helpman, Melitz and Rubinstein, 2008, and Santos Silva and Tenreyro, 2006, 2008). Martin and Pham (2008) discuss some of the relative merits and disadvantages of these approaches in the context of models for trade. They favor the traditional Heckman selection model over two-part models on the grounds that two-part models impose an independence assumption between the decision to trade and the volume of trade, which they consider implausible. There are a number of reasons, however, which incline us to pursue the two-part model in spite of this limitation.

First, there are some reasons concerning convenience. Two-part models compare favorably to the Heckman model regarding computational ease and stability (cf. Martin and Pham, 2008). This is even more important an argument with endogenous regressors such as PTA. While a selection model à la Heckman with common endogenous binary variable has been investigated in econometrics (see Kim, 2006), this model is for an additive error at the level of the outcome equation, and its reformulation to a multiplicative error setup is not trivial as tractability of the model relies partially on the additivity assumption. The two-part model, on the other hand, as seen above, lends itself easily to a generalization with a common endogenous binary variable in both parts. Clearly, these reasons are not substantive, but they should not be downplayed either, especially, in large-scale applications as the present one.[23]

A second reason in favor of two-part models concerns the interpretation of the model. The Heckman selection model was designed for and is superior to the two-part model in applications where potential, unobserved outcomes are of interest, such as reservation wages (cf. Dow and Norton, 2003). This is not the case with trade. Zero trade flows are actual economic outcomes, meaningful 'corner solutions', and not missing data. In the context of trade, it is not evident what potential trade might be, nor how it would be relevant for economic policy. The object of interest here is the effect of PTA on trade flows, both positive and zero, but always actual and observed. The two-part model accomplishes this in a direct and sparse fashion. In contrast to this, Heckman models, while also adaptable to a corner-solution framework, do not model the effect on the intensive margin directly, but rather the effect on potential trade, from which then the intensive margin is recovered.[24]

A third reasons in favor of two-part models concerns performance in estimation. Previous Monte Carlo studies comparing two-part models and Heckman selection models have often found that two-part models outperform selection models in terms of mean squared error even when the data comes from a selection model (see Hay and Olsen, 1984; Manning, Duan, and Rogers, 1987; Leung and Yu, 1996). Martin and Pham (2008) put forward the argument that the performance of the Heckman model could be improved if the correction term's multicollinearity is mitigated by sufficient exogenous variation. We agree. However, it seems very likely that in trade applications multicollinearity will arise inevitably from the necessarily large overlap in the determinants of the decision to trade

[23]In our case, due to importer and exporter fixed effects for 126 countries, a multi-index model as the Heckman or two-part models includes over 500 parameters to be estimated.

[24]Note that the discussion of this point focuses on first-order, bilateral effects.

and of the trading volume.

Last, we return to the independence assumption about the two parts of the model. This might sound stronger than what it really is. It should be stressed that the independence of the two decisions in the two-part model is *conditional on observables*, i.e., – in our specification – after controlling for economic determinants and any importer or exporter specific heterogeneity. In the end, whether selection into trade is endogenous to trade volume is an empirical matter, dependent on data and specification. We resorted to Kim's (2006) model to empirically test the hypothesis of exogenous selection in our model. The formal test failed to reject the conditional independence of the decision to trade from the trading volume.[25]

Therefore, to follow a two-part approach may well be defendend on the grounds of the collective force of the above arguments as well as on the absence of compelling empirical evidence for significant correlation between the parts of the model in our application. To be clear, we are not arguing against the use of selection models in trade in general, nor are we suggesting that two-part models should always be preferred for estimating trade models. Nonetheless, in the present context, it does seem to be the better choice.

6.3 Estimation results

In this subsection, we summarize the parameter estimates from PPML and NLS models described in Section 6.2. Similar to Table 2, Table 4 summarizes the parameter estimates of four alternative models of nominal bilateral exports in U.S. dollars (X_{ij}). Again, every pair of columns gives the parameters of the four covariates of interest in the export out-

[25]Using our specifications from section 6.3 applied to the model of Kim (2006) yielded a t-statistic of 0.18 for a test of exogeneity of the decision to trade. At the same time, a test for exogeneity of PTA was rejected with a t-statistic of 6.66, confirming our previous results. Also, all coefficients exhibited the same signs and significance patterns. As a further check, we estimated Kim's (2006) selection model with endogenous binary regressor again, but this time using a common religion variable as an instrument for the decision to trade (cf. Helpman, Melitz, and Rubinstein, 2008, for rationalizing and using such an instrument). The results did not change substantially. With a t-statistic of 0.30 the conditional independence of the decision to trade from the trading volume could not be rejected, while exogeneity of PTA was still rejected with a t-statistic of 6.48. Detailed results on the selection model estimates are available from the authors upon request.

come equation – PTA_{ij}, DIST_{ij}, BORD_{ij}, and LANG_{ij}. Yet, now we distinguish between the process generating zero versus positive exports and the one generating alternative positive values of exports. The former hurdle process is captured by a probit model for $\mathbb{I}_{ij}$ as explained in Section 6.2, while the latter is estimated via PML. The first pair of columns, columns two to four, gives the parameters with both PPML and NLS when treating PTA_{ij} as exogenous. In the second pair of columns, columns five to seven, we treat PTA_{ij} as endogenous. There, we assume that the processes determining PTA_{ij} and $\mathbb{I}_{ij}$ may be captured by a recursive bivariate probit model for both PPML and NLS.

–– Table 4 ––

Similar to the results in Table 2, we find that the point estimate for PTA_{ij} increases as we abandon the assumption of PTA_{ij} to be exogenous. Again, this result holds true for both PPML and NLS. The point estimates of the one-part and two-part models are relatively similar to each other in broad terms. However, some differences remain, suggesting that zero exports are not generated in sufficient magnitude by the PPML or NLS models. We will provide details on the quantitative impact of PTA membership on the extensive margin of exports in the next section.

Because of this and also due to the non-linear impact of trade frictions or PTA_{ij} on bilateral exports, the corresponding parameter estimate is not as informative of the quantitative importance of PTA membership as in traditional PPML models. But rather, we have to evaluate the role of PTAs for trade flows by means of counterfactual analysis, taking into account third-country effects present in the MR terms in (15) and GDP through equation (13). Such a quantification of the impact of PTA membership on exports in exogenous- versus endogenous-PTA models and a discussion in the light of previous work on the matter is at stake in the subsequent Section 7.

The estimate of ϑ is negative and significant in the PPML model in Table 4, 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 trade when neglecting endogeneity. As a matter of fact, the results in Table 4 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. Such a result could for instance be brought about by a world with sufficiently high market-specific fixed entry costs which are unaffected by PTA formation, whereas marginal delivery costs are lowered by PTA membership. Note that 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. Venturing beyond that point would be purely speculative so that we have to leave it at that.

−− Table 5 −−

Table 5 summarizes the results of the bivariate probit estimation for the PTA equation only. It turns out that the coefficient estimates are very similar to the ones obtained from the univariate probit estimates in Table 3. The parameter estimates of interest which correspond to the extensive margin of exports equation are reported in Table 4.

7 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 2 and 6.1.

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 Abadie and Imbens (2006), for the year 2000; implying an effect of about 97%) 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%) and 0.54 (implying an effect of about 72%). Relative to Baier and Bergstrand's, Magee's (2003) estimated PTA-effects on trade seem extremely large: they amount to in between almost 300 percent and 800 percent! However, these estimates do not account for fixed country effects in both the outcome equation for trade values and the 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, 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 4. 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 is perfectly consistent with a wide range of recent international trade models.[26] Specifically, it captures new trade theory models with love-of-variety preferences and homogenous firms à la Krugman (1980), the Anderson and van Wincoop (2003) exchange economy, the Helpman, Melitz and Rubinstein (2008) model allowing for firm hetero-

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

geneity and zero trade flows, the Eaton and Kortum (2002) Ricardian model, and the Melitz and 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 we apply the Anderson and van Wincoop (2003) framework. Hence, in addition to the parameter estimates, we use the assumption that 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 estimate 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).[27] This can be done by assuming that PTA membership is associated with heterogeneous tariff reductions or not.[28] 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 6a and 6b summarize the predicted effects of PTA formation on trade among PTA members relative to non-members for the eight PPML and NLS models estimated in Tables 2 and 4, respectively.[29] For each model, we distinguish between effects that assume that PTAs alter homogeneous tariffs (in Table 6a) preferentially versus ones that

[27]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 $\sigma = 5$.)

[28]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.

[29]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 thereof.

alter heterogeneous tariffs and (in Table 6b).

−− Tables 6a and 6b −−

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 PPML model which assumes exogenous PTA formation, no specific process for the extensive margin to export, and no heterogeneous tariff effect on trade and GDP in the upper left corner of the table points to an increase in nominal exports among PTA members relative to nonmembers by 54% relative to an equilibrium without any PTAs. This is reflected in the number which is given in the outer left column at the top row of Table 6a labeled "*Average percentage increase of trade flows of members in excess of non-members*". The PPML-based effect is about 45 percentage points higher with endogenous PTA formation (about 99% higher exports among PTA members relative to nonmembers than without PTAs; see the results in the third column at the top of Table 6a). Ignoring the heterogeneity of tariffs brings about a negligible bias in our application.[30] To see the latter, compare the results at the top of Table 6a with the corresponding ones in Table 6b.

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 preferred PPML model in the seventh column of Table 6a by about 11 percentage points from about 99% to 110%. Of the average PTA-induced effect on exports of 110%, about 10 percent are contributed by the extensive margin while the rest is due to the intensive margin. In general, the contribution to the average PTA-induced effect on exports of the extensive margin in the two-part models can be inferred by comparing the difference between the corresponding one-part and two-part models.[31] Overall, the estimated (long-run) effects of PTA membership on bilateral trade are quite large.

[30]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 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 a lot of information.

[31]Of the 9,891 country-pairs with predicted positive bilateral exports in the cum-PTA benchmark

Moreover, Tables 6a and 6b 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.[32] To see this, consider the two blocs of results in the lower parts of Tables 6a and 6b. 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.[33] 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

equilibrium, 177 would stop exporting if all PTAs were abandoned. This result is based on estimates which disregard the fact that (most-favored nation) tariffs are heterogeneous across countries so that preferential 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.

[32]That treatment effects tend to be heterogeneous across the treated is widely acknowledged in other fields of economics (see Bitler, Gelbach, and Hoynes, 2005, for an example in public economics). However, international economists tend to focus on average effects of treatments such as PTA membership but also other treatments on outcome of interest and tend to ignore that theoretical models often would suggest heterogeneous treatment effects.

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

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 micro-econometric methods (such as propensity score matching or switching regression) did not share this feature.

8 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 continuous trade policy measures – to be endogenous; and they account for non-linear 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 data-set 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 more than 40 percentage points higher than in a model which assumes PTA membership to be exogenous. With the data-set 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 11 percentage points as compared to the preferred model.

References

Abadie, Alberto and Guido W. Imbens (2006), Large sample properties of matching estimators for average treatment effects, *Econometrica* **74**, 235–267.

Aitken, Norman D. (1973), The effect of the EEC and EFTA on European trade: A temporal cross-section analysis, *American Economic Review* **63**, 881-892.

Anderson, James E. (1979), A theoretical foundation for the gravity equation, *American Economic Review* **69**, 106-116.

Anderson, James E. (2009), Gravity, productivity and the pattern of production and trade, *NBER Working Paper No. 14642*.

Anderson, James E. and Eric van Wincoop (2002), Borders, trade, and welfare, in: Susan M. Collins and Dani Rodrik (eds.), *Brookings Trade Forum 2001*, Washington: Brookings Institution, 207-244.

Anderson, James E. and Eric van Wincoop (2003), Gravity with gravitas: A solution to the border puzzle, *American Economic Review* **93**, 170-192.

Baier, Scott L. and Jeffey H. Bergstrand (2002), On the endogeneity of international trade flows and free trade agreements, unpublished manuscript, University of Notre Dame.

Baier, Scott L. and Jeffrey H. Bergstrand (2004), Economic determinants of free trade agreements, *Journal of International Economics* **64**, 29-63.

Baier, Scott L. and Jeffey H. Bergstrand (2007), Do free trade agreements actually increase members' international trade?, *Journal of International Economics* **71**, 72-95.

Baier, Scott L. and Jeffey H. Bergstrand (2009), Estimating the effects of free trade agreements on international trade flows using matching econometrics, *Journal of International Economics* **77**, 63–76.

Baldwin, Richard E. (2008), Big-think regionalism: a critical survey, *NBER Working Paper No. 14056*.

Baldwin, Richard E. and Anthony J. Venables (1995), Regional economic integration, in: Gene M. Grossman and Kenneth R. Rogoff (eds.), *Handbook of International Economics, Volume 3*, chapter 31, Amsterdam: Elsevier Science, 1597-1644.

Bitler, Marianne, Jonah Geldbach, and Hilary W. Hoynes (2006), What mean impacts miss: Distributional effects of welfare reform experiments, *American Economic Review* **96**, 988-1012.

Cameron, Colin A. and Pravin K. Trivedi (2005), *Microeconometrics. Methods and Applications*, University Press: Cambridge, MA.

Carrère, Céline (2006), Revisiting the effects of regional trade agreements on trade flows with proper specification of the gravity model, *European Economic Review* **50**, 223-247.

Cragg, John G. (1971), Some statistical models for limited dependent variables with application to the demand for durable goods, *Econometrica* **39**, 829-844.

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**, 5-18.

Duan, Naihua, Willard G. Manning, Carl N. Morris, and Joseph P. Newhouse (1983), A comparison of alternative models for the demand for medical care, *Journal of Business and Economic Statistics* **1**, 115-126.

Dixit, Avinash K. and Joseph E. Stiglitz (1977), Monopolistic competition and optimum product diversity, *American Economic Review* **67**, 297-308.

Eaton, Jonathan and Samuel Kortum (2002), Technology, geography, and trade, *Econometrica* **70**, 1741-1779.

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**, 478-510.

Egger, Hartmut, Peter H. Egger, and David Greenaway (2008), Trade liberalization with multinational firms: Effects on welfare and intra-industry trade, *Journal of International Economics* **74**, 278-298.

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

Felbermayr, Gabriel and Wilhelm Kohler (2006), Exploring the intensive and extensive margins of world trade, *Review of World Economics* **142**, 642-674.

Freund, Caroline (2000), Different paths to free trade: The gains from regionalism, *Quarterly Journal of Economics* **115**, 1317-1341.

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 edition, Prentice Hall: New Jersey.

Greenaway, David and Chris Milner (2002), Regionalism and gravity, *Scottish Journal of Political Economy* **49**, 574-585.

Hay, Joel W. and Randall J. Olsen (1984), Let Them Eat Cake: A Note on Comparing Alternative Models of the Demand for Medical Care, *Journal of Business and Economics Statistics* **2**, 279-282.

Heckman James J., Lance Lochner, and Christopher Taber (1998), General-equilibrium treatment effects: A study of tuition policy, *American Economic Review* **88**, 381-386.

Helpman Elhanan, Marc Melitz, and Yona Rubinstein (2008), Estimating trade flows: Trading partners and trading volumes, *Quarterly Journal of Economics* **2**, 441-487.

Kim, Kyoo il (2006), Sample selection models with a common dummy endogenous regressor in simultaneous equations: A simple two-step estimation, *Economics Letters 91*, 280-286.

Krugman, Paul R. (1980), Scale economies, product differentiation, and the pattern of trade, *American Economic Review* **70**, 950-959.

Leung, Siu Fai and Shihti Yu (1996), On the choice between sample selection and two-part models, *Journal of Econometrics* **72**, 197-229.

Magee, Christopher S. (2003), Endogenous preferential trade agreements: An empirical analysis, *Contributions to Economic Analysis & Policy* **2**, Article 15.

Manning, Willard G., Naihua Duan and William H. Rogers (1987), Monte Carlo Evidence on the choice between sample selection and two-part models, *Journal of Econometrics* **35**, 59-82.

Marshall, Monty G. and Keith Jaggers (2007), Polity IV project: Political regime characteristics and transitions, 1800-2006, Center for International Development and Conflict Management, University of Maryland.

Martin, Will J. and Cong S. Pham (2008), Estimating the Gravity Model When Zero Trade Flows Are Frequent, Economics Series 2008 03, Deakin University, Faculty of Business and Law, School of Accounting, Economics and Finance.

Melitz, Marc and Gianmarco I.P. Ottaviano (2008), Market size, trade, and productivity, *Review of Economic Studies* **75**, 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**, 271-282.

Rodrik, Dani (1995), Political economy of trade policy, in: G.M. Grossman and K.R. Rogoff (eds.), *Handbook of International Economics, Volume 3*, Elsevier Science: Amsterdam, chapter 28, 1457-1494.

Santos Silva, João M.C. and Silvana Tenreyro (2006), The log of gravity, *Review of Economics and Statistics* **88**, 641-658.

Santos Silva, João M.C. and Silvana Tenreyro (2008), Trading partners and trading volumes: Implementing the Helpman-Melitz-Rubinstein model empirically, Discussion Paper No. 662, Department of Economics, University of Essex.

Soloaga, Isidro and L. Alan Winters (2001), Regionalism in the nineties: What effect on trade?, *North American Journal of Economics and Finance* **12**, 1-29.

Terza, Joseph V. (1998), Estimating count data models with endogenous switching: Sample selection and endogenous treatment effects, *Journal of Econometrics* **88**, 129-154.

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

Windmeijer, Frank A.G. and João M.C. Santos Silva (1997), Endogeneity in count data models: An application to demand for health care, *Journal of Applied Econometrics* **12**, 281-294.

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

Appendix

A Country coverage (126 economies)

The following set of countries is covered in our data-set:

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.

B PTA coverage (121 agreements)

Our data-set 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 data-set:

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).

Tables

Table 1: Descriptive statistics of the data

Variable	Description	Mean	Std.Dev.	Min.	Max.
X_{ij}	nominal exports in million U.S. dollars	305.9274	3257.2670	0	213763.06
$\mathbb{I}_{ij}$	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 i and j and the 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
Number of observations			15750		

Table 2: Estimation results for structural gravity models for trade

	Exogenous PTA			Endogenous PTA	
	OLS	PPML	NLS	PPML	NLS
Dep. var.:	$\ln(X_{ij})$	X_{ij}	X_{ij}	X_{ij}	X_{ij}
(1)	(2)	(3)	(4)	(5)	(6)
PTA_{ij}	0.3156	0.4855	0.6247	0.7679	0.8880
	(0.0448)	(0.0623)	(0.1000)	(0.1257)	(0.2068)
DIST_{ij}	-1.3572	-0.7000	-0.6248	-0.6278	-0.5581
	(0.0292)	(0.0294)	(0.0397)	(0.0416)	(0.0569)
BORD_{ij}	0.6306	0.6585	0.5923	0.6555	0.5639
	(0.1063)	(0.0634)	(0.0755)	(0.0622)	(0.0713)
LANG_{ij}	0.7186	0.2197	0.2534	0.2498	0.2862
	(0.0539)	(0.0646)	(0.0816)	(0.0622)	(0.0799)
$\hat{\vartheta}$	–	–	–	-0.1963	-0.1776
				(0.0853)	(0.1258)
Number of observations	9891	15750	15750	15750	15750
Number of countries	126	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. Columns (5) and (6) indicate results where PTA was instrumented. $\hat{\vartheta}$ is a measure for potential endogeneity of PTA_{ij}. PTA_{ij}, BORD_{ij} (common border), and LANG_{ij} (common language/ethnicity) are binary variables. Robust, two-step adjusted standard errors in parenthesis.

Table 3: Probit estimation results for reduced form equation for PTA

Dep. var.: PTA	Coeff.	Std.err.
DURAB_{ij}	0.0059	0.0030
DURAB^2_{ij}	-0.0001	0.0000
POLCOMP_{ij}	-0.1083	0.0133
POLCOMP^2_{ij}	0.0011	0.0001
AUTOC_{ij}	0.0737	0.0155
AUTOC^2_{ij}	-0.0007	0.0002
DIST_{ij}	0.2332	0.3715
DIST^2_{ij}	-0.0812	0.0245
DKL_{ij}	-0.2251	0.0188
DROWKL_{ij}	0.9126	0.3203
RGDPsum_{ij}	-0.0268	0.1448
RGDPsim_{ij}	0.0609	0.0747
LANG_{ij}	-0.0909	0.0645
CONT_{ij}	0.6852	0.0501
COLONY_{ij}	-0.0012	0.1953
COMCOL_{ij}	0.4935	0.0736
CURCOL_{ij}	0.8001	0.2538
SMCTRY_{ij}	1.3177	0.2516
Number of observations	15750	
Number of countries	126	

Notes:
Data source: United Nations' World Trade Database, World Bank's World Development Indicators.
Further control variables include importer and exporter fixed effects and a constant.

Table 4: Estimation results for two-part gravity models for trade

	Exogenous PTA			Endogenous PTA		
	$\Pr(\mathbb{I}_{ij}=1\|\cdot)$	$\mathrm{E}(X_{ij}\|.,\mathbb{I}_{ij}=1)$		$\Pr(\mathbb{I}_{ij}=1\|\cdot)$	$\mathrm{E}(X_{ij}\|.,\mathbb{I}_{ij}=1)$	
	Probit	PPML	NLS	Biv. Probit	PPML	NLS
(1)	(2)	(3)	(4)	(5)	(6)	(7)
PTA_{ij}	0.3647	0.4789	0.6215	-0.0879	0.7690	0.8789
	(0.0555)	(0.0626)	(0.1001)	(0.1116)	(0.1243)	(0.2049)
DIST_{ij}	-1.1950	-0.7023	0.5921	-1.2736	-0.6269	-0.5603
	(0.0384)	(0.0296)	(0.0755)	(0.0410)	(0.0417)	(0.0569)
BORD_{ij}	-0.4388	0.6589	-0.6262	-0.3120	0.6563	0.5642
	(0.1681)	(0.0632)	(0.0398)	(0.1679)	(0.0619)	(0.0714)
LANG_{ij}	0.6415	0.2164	0.2531	0.6180	0.2464	0.2846
	(0.0629)	(0.0644)	(0.0815)	(0.0627)	(0.0620)	(0.0801)
$\hat{\rho}_{v\xi}$	–	–	–	0.3012	–	–
				(0.0637)		
$\hat{\vartheta}$	–	–	–	–	-0.2028	-0.1754
				–	(0.0843)	(0.1271)
Number of observations	13500	9891	9891	15750	9891	9891
Number of countries	126	126	126	126	126	126

Notes:
All regressions include importer and exporter fixed effects.
In column (2), 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. Standard errors in parenthesis. Standard errors for $\mathrm{E}(X_{ij}|.,\mathbb{I}_{ij}=1)$ are robust and two-step adjusted.

Table 5: Bivariate probit estimation results for PTA

Dep. var.: PTA	Coeff.	Std.err.
DURAB_{ij}	0.0066	0.0029
DURAB_{ij}^2	-0.0001	0.00003
POLCOMP_{ij}	-0.1068	0.0132
POLCOMP_{ij}^2	0.0010	0.0001
AUTOC_{ij}	0.0794	0.0151
AUTOC_{ij}^2	-0.0007	0.0001
DIST_{ij}	-0.5767	0.2717
DIST_{ij}^2	-0.0303	0.0184
DKL_{ij}	-0.2206	0.0188
DROWKL_{ij}	0.9889	0.3091
RGDPsum_{ij}	-0.0083	0.1318
RGDPsim_{ij}	0.0700	0.0682
LANG_{ij}	-0.0560	0.0637
CONT_{ij}	0.6867	0.0501
COLONY_{ij}	-0.0195	0.1907
COMCOL_{ij}	0.3855	0.0756
CURCOL_{ij}	0.8021	0.2412
SMCTRY_{ij}	1.0335	0.2191
Number of observations	15750	
Number of countries	126	

Notes:
Data source: United Nations' World Trade Database, World Bank's World Development Indicators.
Further control variables include importer and exporter fixed effects and a constant.
The second equation of the bivariate probit model is a regression of the probability to trade on PTA, log distance, common border, common language/ethnicity, importer and exporter fixed effects and a constant. Results are reported in Table 4

Table 6a: Counterfactual results, homogenous tariff rates

	One-part models				Two-part models[b]			
	Exogenous PTA		Endogenous PTA		Exogenous PTA		Endogenous PTA	
	PPML	NLS	PPML	NLS	PPML	NLS	PPML	NLS
Average percentage increase of trade flows of PTA members in excess of non-members	54.27	74.33	98.66	120.10	63.01	83.98	110.07	129.45
ΔX_{ij} among PTA members in %:[a]								
mean	20.0853	27.4616	34.3624	42.1099	18.3858	25.3532	31.9194	38.4637
std. dev.	20.8203	28.7353	37.1613	45.9208	19.1919	26.7205	34.7583	42.1930
min	-31.0661	-38.0643	-45.2417	-50.1009	-30.7348	-37.9412	-45.3082	-49.7424
max	89.0851	128.2430	174.0078	222.9863	82.1443	115.3547	156.3760	189.3391
# of PTA member pairs with								
positive effect	2734	2735	2710	2713	2784	2781	2764	2761
negative effect	240	239	264	261	240	243	260	263
ΔX_{ij} among PTA non-members in %:[a]								
mean	-8.4788	-10.4377	-12.8658	-14.2747	-8.5206	-10.5419	-13.0371	-14.3023
std. dev.	11.0091	13.4688	16.1528	17.8867	10.9200	13.4392	16.1881	17.7507
min	-53.9526	-62.8453	-71.2951	-76.0053	-53.4747	-62.6759	-71.3592	-75.6461
max	13.5321	18.6218	21.9596	27.1705	11.6169	14.6234	17.9313	19.9144
# of PTA non-member pairs with								
positive effect	648	637	585	576	566	560	511	513
negative effect	6224	6235	6287	6296	6256	6262	6311	6309

Notes:

[a] ΔX_{ij}: base scenario trade flows minus counterfactual trade flows relative to counterfactual trade flows in %; 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.

There are 3,609 PTA member pairs, whereof in 635 (585) $\Delta X_{ij} = 0$ occurred in the one-part (two-part) models. Of the 12,016 PTA non-member pairs, in 5,144 (5,194) cases the model predicted $\Delta X_{ij} = 0$.

Table 6b: Counterfactual results, heterogenous tariff rates

	One-part models				Two-part models[b]			
	Exogenous PTA		Endogenous PTA		Exogenous PTA		Endogenous PTA	
	PPML	NLS	PPML	NLS	PPML	NLS	PPML	NLS
Average percentage increase of trade flows of PTA members in excess of non-members	54.17	74.19	98.52	119.93	62.87	83.80	109.88	129.24
ΔX_{ij} among PTA members in %:[a]								
mean	19.7942	27.0693	33.9699	41.6003	18.0771	24.9497	31.5141	37.9557
std. dev.	20.9280	28.7512	37.1605	45.8032	19.3134	26.7589	34.7823	42.1189
min	-31.6092	-38.3996	-45.5465	-50.4034	-31.3305	-38.2633	-45.6025	-50.0361
max	87.9299	126.3383	171.8040	219.7474	79.8197	112.5929	153.5508	188.1116
# of PTA member pairs with								
positive effect	2726	2723	2710	2715	2777	2774	2761	2759
negative effect	248	251	264	259	247	250	263	265
ΔX_{ij} among PTA non-members in %:[a]								
mean	-8.5979	-10.5893	-13.0031	-14.4396	-8.6385	-10.6909	-13.1720	-14.4632
std. dev.	11.1147	13.5771	16.2462	17.9829	11.0350	13.5559	16.2894	17.8548
min	-54.1725	-63.0465	-71.4552	-76.1490	-53.7025	-62.8824	-71.5227	-75.7942
max	13.3921	18.1760	21.5421	26.4472	11.0525	13.9861	17.2845	19.2018
# of PTA non-member pairs with								
positive effect	637	624	574	576	561	550	500	510
negative effect	6235	6248	6298	6296	6261	6272	6322	6312

Notes:

[a] ΔX_{ij}: base scenario trade flows minus counterfactual trade flows relative to counterfactual trade flows in %; 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.

There are 3,609 PTA member pairs, whereof in 635 (585) $\Delta X_{ij} = 0$ occurred in the one-part (two-part) models. Of the 12,016 PTA non-member pairs, in 5,144 (5,194) cases the model predicted $\Delta X_{ij} = 0$.