

Interdependent Preferential Trade Agreement Memberships: An Empirical Analysis*

Peter Egger[†] and Mario Larch[‡]

August 1, 2006

Abstract

Recent theoretical work on bilateral trade preferences stresses their dependence on but also their consequences for the multilateral trading system. In particular, a country's choice of participating in a preferential trade agreement (PTA) depends on the choice of other economies to participate therein. However, recent empirical work on the determinants of PTA formation assumes that countries are independent in that regard. This paper lays out an empirical analysis to study the role of interdependencies in PTA membership in a large data-set of 15,753 country-pairs. Applying modern econometric techniques, a PTA membership is found to create an incentive for other country-pairs to participate in a PTA as well. Especially, countries have an incentive to participate in the same PTA if their neighbors are members already.

Key words: Preferential trade agreements; Limited dependent variable models; Spatial econometrics

JEL classification: F14; F15; C11; C15; C25

* *Acknowledgements:* To be added.

[†] *Affiliation:* Ifo Institute for Economic Research, Ludwig-Maximilian University of Munich, CESifo, and Centre for Globalization and Economic Policy, University of Nottingham. *Address:* Ifo Institute for Economic Research, Poschingerstr. 5, 81679 Munich, Germany.

[‡] *Affiliation:* Ifo-Institute. *Address:* Poschingerstr. 5, 81679 Munich, Germany.

1 Introduction

If everything in the universe depends on everything in a fundamental way, it might be impossible to get close to a full solution by investigating parts of the problem in isolation.

(Stephen Hawking and Leonard Mlodinov, 2005, *A Briefer History of Time*, Bantam Dell, New York, p. 15)

The continued integration of the European Union (EU), the formation of the North American Free Trade Agreement (NAFTA) and the membership of Mexico therein, as well as the political discussion about the formation of a preferential trade agreement (PTA) between the Americas have been major sources for the renewed interest in PTAs in the last two decades. With the increasing globalization of the world economy, it seems that there is a raising concern about the global consequences of regionalism. For instance, Bagwell and Staiger (1997a,b, 1999, 2005) find that PTAs form obstacles to the delivery of efficient multilateral trade policy under the auspices of the General Agreement on Trade and Tariffs (GATT) and the World Trade Organization (WTO). A similar conclusion is reached in Bond, Riezman, and Syropoulos (2004). While this line of research mainly focuses on the interference of PTAs with an efficient multilateral trade liberalization, another sub-literature is concerned with the interdependence of regionalism and PTA memberships as such. Three main reasons for an interdependence of PTA memberships have been put forward in previous work: political-economy forces, optimal bloc size and the related market power, and economic and geographical fundamentals.

Political-economy forces are crucial in the so-called *domino theory of regionalism* introduced by Baldwin (1995, 1997). There, countries desire to participate in an existing PTA since the threat of a loss in the export sector associated with non-participation nourishes lobbying activity to promote such a participation. If no such accession is feasible for political reasons such countries might prefer engaging in a new PTA with other outsiders for similar reasons. The establishment of both NAFTA and the European Single

Market created tremendous asymmetries among firms with and without access to these huge markets. Market access is particularly important in a world where firms are mobile across borders and multinationals control large part of the goods trade. Then, market integration through trade liberalization creates an incentive for multinational plant location and stimulates a capital influx from abroad (see Baldwin, Forslid, and Haaland, 1996, for simulation-based evidence; Yeaple, 2003, and Helpman, Grossman, and Szeidl, 2006, for models with complex multinationals; and Ekholm, Forslid, and Markusen, 2003, for a theoretical model of export-platform multinationals). In turn, the threat of capital flight into PTAs exerts a pressure on outsiders to join existing PTAs or to found new ones.¹ This view is largely supported, for instance, by Abbott (1999) who argues in Chapter III-C of his monograph on the North American integration regime that "*the NAFTA was in part negotiated to counterbalance the growing economic and political influence of the EU. The EU has since pursued negotiations with Mercosur and with Mexico on closer economic relations.*" At the same time PTAs were negotiated in Asia and Oceania. And the Central and Eastern European economies have not only reduced their trade barriers bilaterally with the members of the EU and the European Free Trade Area (EFTA) but most of them have become EU members themselves in 2004.

Issues with optimal bloc size are analyzed by Krugman (1991a) who elaborates on the consequences of free trade within an arbitrary number of fully symmetric PTAs in the world economy. In this setting, it is bad for world welfare if there are only a few, large PTAs that exert market power by setting high external tariffs. But rather, it is desirable to either have a single PTA or a large number of PTAs, where external tariffs are low in equilibrium. Bond and Syropoulos (1996) extend this framework by exploring the role of the size of asymmetric blocs for optimum external tariffs (i.e., market power) and both bloc and world welfare.

Economic and geographical fundamentals comprise the relative size of countries, their relative factor endowments, and trade costs. Frankel, Stein, and Wei (1995) set up a new

¹Whalley (1996) puts forward a different reason for PTA membership, namely the threat of economies of standing alone as PTA outsiders during trade wars.

trade theory model of many countries that are grouped into continents (with high trade costs across continents and low ones within them). They point out that PTAs should be formed along natural continental lines but, from a world welfare perspective, the PTAs should be partial. They conclude that "*the world trading system is currently in danger of entering the zone of excessive regionalization*" (ibid., p. 92). The WTO should promote extending the scope of existing PTAs rather than fully eliminate intra-PTA tariffs, which would imply a departure from the principles of GATT expressed in its Article XXIV. Baier and Bergtrand (2004, henceforth BB) follow Krugman (1991b) closely by focusing on three continents with two economies each. Their goal is to motivate an empirical model of endogenous selection into PTAs depending on intra- and intercontinental trade costs, country size, and relative factor endowment differences. They confirm Bhagwati's (1993) and Krishna's (2003) view that positive welfare effects of PTAs are more likely for countries that already trade disproportionately with each other. In particular, BB's hypotheses are that (i) countries with lower bilateral trade costs, (ii) ones with higher trade costs from the rest of the world, (iii) countries with a greater similarity in country size and relative factor endowments, (iv) and ones that are relatively dissimilar in these regards from the rest of the world are expected to face high welfare gains from entering a PTA. However, while a country's PTA membership affects the other economies' welfare in the model of BB, it was not their task to shed light on this issue explicitly.

This paper lays out an empirical analysis of the determinants of PTA membership by explicitly accounting for the interdependence of membership events. In particular, we extend the analysis of BB in several directions. First, we illustrate how the welfare effects of a PTA formation in their model depends on whether other country-pairs form PTAs as well. In this regard, we distinguish between a country's incentive to join a PTA with several other economies versus forming one with other outsiders. Higher positive welfare effects of a PTA at the bilateral level are then associated with a greater empirical probability of participating in a PTA. Second, we set up an empirical model where the probability of a PTA membership depends on economic and geographical fundamentals plus the probability of other country-pairs to participate in the same PTA or other ones.

Consistent estimation requires recent econometric techniques for interdependent limited-dependent variable problems. Third, we put forward empirical results that are based on data covering 15,753 country-pairs – which is more than ten times the sample size in the study of BB.

Our empirical findings regarding the economic fundamentals largely support the ones put forward in BB in this much larger sample of country-pairs. Beyond that, there is a strong interdependence of PTA memberships that declines in geographical distance being in line with the hypotheses central to this paper. In particular, this interdependence is stronger for PTA memberships in one and the same PTA. Hence, a PTA membership increases the probability of countries to join an adjacent PTA, and it increases the probability of outsiders to integrate in other PTAs, but to a somewhat lesser extent. As will be illustrated below, interdependence is also found in the sample of country-pairs covered by BB. And the findings are largely robust to the omission of political variables and to variations in the decay of interdependence through geographical distance.

The remainder of the paper is organized as follows. In the next section we reconcile the model by BB to put forward three hypotheses regarding the PTA-related interdependence of country-pairs. Section 3 lays out the empirical model that accounts for interdependent observations with limited dependent variables. Section 4 summarizes the estimation results for our data-set of 15,753 country-pairs, and Section 5 assesses their robustness with respect to alternative specifications. The last section concludes with a short summary of the most important findings.

2 Interdependence in the Baier and Bergstrand (2004) model

BB formulate a numerically solvable variant of the model by Krugman (1991b) and Frankel, Stein, and Wei (1995, 1998) to derive their set of empirically testable hypotheses. It is not our goal to elaborate on a new theoretical framework here. But rather, we intend

to deduce three PTA-related hypotheses regarding the interdependence of country-pairs' welfare from the model of BB that have not been spelled out and tested there. To keep matters as simple as possible, in the sequel we use the same parameter values as reported in BB and focus on the welfare effects of PTAs on their members.

The model of BB consists of three continents (A, B, C) with two countries each ($1A, 2A, 1B, 2B, 1C, 2C$). There are two sectors, manufacturing (G) and services (S), where each sub-utility is characterized by Dixit and Stiglitz (1977) type preferences and a love of variety of consumers. Income consists of the factor income of the two primary production factors, labor (L) and capital (K), plus external tariff revenues collected from imports entering from outside a PTA. Trade is not only impeded by ad-valorem tariffs but also by non-tariff iceberg trade costs. In this framework, intercontinental iceberg trade costs may differ from intracontinental ones. The total costs of serving consumers on other continents include intracontinental trade costs plus the intercontinental ones. Hence, as in BB, in the subsequent analysis intercontinental trade costs are to be interpreted as the additional costs of serving consumers intercontinentally rather than intracontinentally. We suppress a formal exposition of the model in the main text but relegate this to Appendix A to facilitate the reading.

One of the central findings in BB is the role of intra- versus intercontinental trade costs for the net welfare effects of intra- versus intercontinental PTAs.² As a starting point for our analysis, let us reproduce Figure 1 of BB. There, the underlying assumption is that the three continents are populated with two identical economies each.

— Figure 1 —

There are two surfaces displayed in the figure, one relates to *natural* (i.e., intracontinental) and a second one to *unnatural* (i.e., intercontinental) PTAs. Independent of the size of intercontinental trade costs, the net welfare gain of a natural PTA is always at

²Other important insights in their work relate to the interaction of country size or relative factor endowment differences and intra- versus intercontinental trade costs in triggering the welfare effects of trade liberalization. It is beyond the scope of this paper to go into detail regarding country size and factor endowment issues. In this regard, our empirical analysis rests on the hypotheses put forward in BB's work.

least as high as that one of an unnatural PTA in the figure. Hence, the surface for natural PTAs is always on top of that one for unnatural PTAs.

However, what is not obvious from the figure is that – in either case, natural and unnatural PTA formation – three PTAs are implemented simultaneously. Hence, with natural PTAs all three continents eliminate their intracontinental tariffs, keeping intercontinental ones at their original level. And with unnatural PTAs there are three cross-continental country-pairs that implement non-overlapping PTAs at the same time (e.g., $1A$ with $1B$, $2A$ with $1C$, and $2B$ with $2C$). This implies that the effects displayed in Figure 1 consist of two components. On the one hand, there are welfare gains for each joining country due to the net creation of trade with the PTA they are involved in. On the other hand, there is a welfare loss due to a redirection of trade induced by foreign PTAs. Our goal is to disentangle these effects to derive three hypotheses relating to the interdependence of PTAs. These hypotheses will be at the heart of our empirical analysis.

2.1 Interdependence 1: Net welfare effects of a PTA if foreign PTAs are implemented

As indicated before, Figure 1 sheds light on how a country's net welfare changes due to the formation of three PTAs simultaneously as compared to a situation without any PTA worldwide. However, to study the role of interdependence, we need to rely on a different counterfactual, namely one where, e.g., natural PTAs are formed on continents B and C but not on A . The net welfare gain of forming a PTA for an economy on continent A , say country $1A$, is then the difference between the original natural PTA scenario as in Figure 1 and the counterfactual where only countries $1A$ and $2A$ do not establish a PTA. Similarly, we may conduct such a thought experiment for unnatural PTAs. For instance, compare the unnatural trade liberalization scenario with three PTAs in BB with one where countries $2A$ and $1C$ and countries $2B$ and $2C$ liberalize trade but countries $1A$ and $1B$ do not. The corresponding results for both natural (on top) and unnatural PTAs (at the bottom) are displayed in Figure 2.

Let us start with the discussion of natural PTAs in Figure 2. Obviously, there is an unambiguous net welfare gain from implementing a PTA given that the foreign economies do so as well. A comparison of Figures 2 and 1 indicates that the welfare gain of forming a PTA given that foreign countries do so is even larger than without foreign PTAs. To see this, let us subtract the surface in Figure 2 from the one in Figure 1. This results in negative values throughout, indicating that an outsider faces a net welfare loss induced by foreign PTA formation. The reason behind this result is Baldwin's (1995) domino effect.³ The countries at continent *A* lose on net if natural PTAs are formed at continents *B* and *C* and, as a consequence, trade is diverted away from *A*. The net welfare gain from PTAs in Figure 2 increases almost exponentially as intracontinental trade costs decline.

A qualitatively similar result is obtained for unnatural PTAs. But the corresponding net welfare gains are at most as high as those of natural PTAs. Note that a simultaneous formation of three unnatural PTAs leads to non-positive net welfare effects in Figure 1. However, Figure 2 illustrates that it is still better for a country to establish an unnatural PTA than none, given that there are foreign unnatural PTAs. But the incentive to do so is lower in welfare terms than for a natural PTA. The net welfare gain of forming an unnatural PTA rises exponentially if inter- and/or intracontinental trade costs decline. Note that net welfare from unnatural PTA formation responds with a similar sensitivity to a decline in intracontinental trade costs as it does with respect to intercontinental ones. The reason is that the redirection of trade arising from its natural trade partner's PTA with an unnatural trade partner is large if both intra- and intercontinental trade costs are low.

³In fact, Baldwin's argument is richer, since he also accounts for the role of mobile production capital. But this is beyond the scope of the analysis, here.

2.2 Interdependence 2: Welfare effects from joining others in forming a PTA

We have seen from the previous discussion that countries have an incentive to form PTAs in particular if there are other PTAs in place. And this incentive is always at least as large for a natural PTA as for an unnatural one. However, so far each country was either a member or an outsider of a two-country PTA. Hence, there was no possibility for, say, economy $1A$ to join another country-pair and form a three-country PTA. It is this section's task to consider that option. In particular, we are interested in how a country's welfare may be differently affected when joining another country-pair as compared to forming a separate PTA with another outsider.⁴

To shed light on this issue, consider a liberalization scenario where country $1A$ joins countries $2A$ and $1C$. At the same time, $2B$ forms a PTA with $2C$ so that only $1B$ is left outside. Let us again compare the associated net welfare effect on country $1A$ with the scenario where three unnatural PTAs are formed as in the bottom surface of Figure 1.⁵ The differential effect is displayed in Figure 3.

— Figure 3 —

Obviously, the net welfare gain is higher for country $1A$ if it joins two countries in an unnatural PTA than from forming an unnatural two-country PTA with an outsider. The reason is that the trade creation effects are higher in the former case. This is enforced by but not fully attributable to the fact that one of the two bloc members is located at the same continent. The difference in the net welfare gain from joining a PTA to the baseline scenario with three unnatural PTAs is non-monotonous in intercontinental trade costs. Recall the surfaces displaying a country's net welfare gain from forming a natural or an unnatural PTA (when two other pairs do so as well) relative to a counterfactual scenario

⁴Since our contribution is primarily empirical, it is beyond this paper's scope to consider all possible permutations of bloc size or alternative numbers of countries and/or continents. In this regard, we steer the interested reader to the work of Bond and Syropoulos (1996) and Frankel, Stein, and Wei (1995).

⁵Recall that the three unnatural PTAs in Figure 1 are: $1A$ with $1B$, $2A$ with $1C$, and $2B$ with $2C$. Notice that at zero intercontinental trade costs any differences between intra- and intercontinental relationships are eliminated in the BB model.

without any PTAs from Figure 1. Figure 3 illustrates that joining a natural and an unnatural trade partner in a PTA leads to both trade creation and trade redirection inside the PTA. Higher intercontinental trade costs redirect trade from the unnatural PTA member to the natural one, leading to welfare gains for the two natural partners. However, if the intercontinental trade costs become too high, the associated trade destruction through redirection may outweigh the welfare gains from trade creation even for the natural trade partners. A reduction in intracontinental trade costs leads to a reduction in iceberg trade costs for all trading partners.⁶ Therefore, the difference in the net welfare gain in Figure 3 declines monotonously in intracontinental trade costs.

— Figure 4 —

In Figure 4, we undertake an alternative experiment, where the same country 1A joins a foreign natural PTA rather than an unnatural one as in Figure 3. Notice that the relationship between intercontinental trade costs and the joining country's net welfare is now monotonous. For instance, the country may gain less than in a natural PTA with country 2A if intracontinental trade costs are low but intercontinental ones are high. Then, trade with the foreign PTA is mainly constrained by intercontinental iceberg trade costs and the response of both the trade volume and the net welfare to eliminating tariffs with two unnatural trade partners is small. Accordingly, the loss from not eliminating the tariff barriers with natural trade partner is big. Hence, the country would do better by staying outside the foreign PTA and form one with its natural trade partner. The net welfare gain (loss) declines (increases) as intra- and intercontinental trade costs increase proportionately. The previous analysis supports three hypotheses that may be investigated in the subsequent empirical investigation.

Figures 3 and 4 suggest that joining a PTA is particularly beneficial if average trade costs with the PTA members are low. For instance, joining an unnatural PTA as a natural trade partner of one of the member countries in Figure 3 is particularly beneficial

⁶Recall that BB formulate trade costs such that the total costs of serving consumers on other continents are made up of both intra- and intercontinental trade costs.

if intracontinental trade costs are low. Entering a foreign natural PTA as an unnatural trade partner in Figure 4 increases a country's net welfare by more if intercontinental trade costs are low.

2.3 Summary of new empirical hypotheses

We may summarize the three hypotheses relating to interdependence in PTA memberships in the following way.

Hypothesis 1: Foreign PTAs increase the incentive for a country to form or enter a PTA, irrespective of whether natural or unnatural PTAs are concerned. Hence, this hypothesis is about interdependence of PTA memberships as such.

Hypothesis 2: This interdependence declines in trade costs. Although this hypothesis cannot be inferred separately from Hypothesis 1, it is explicit about the channel through which the interdependence works, namely trade costs.

Hypothesis 3: The formation of a PTA creates an incentive for outsiders to join – in particular at low trade costs – that is stronger than the one to create or join another PTA in response. Hence, interdependence is stronger within than across PTAs.

3 An empirical approach to interdependent PTA memberships

3.1 The problem: cross-sectional interdependence with binary variables

Empirical applications treat PTA membership as a binary variable with entry one if two countries are members of the same PTA and zero else (see Magee, 2003, and BB). The binary outcome of PTA participation may be viewed as a reflection of the difference in unobservable utility between membership and non-membership scenarios as in the BB model (see McFadden, 1974, and Domencich and McFadden, 1975, for a random utility

interpretation of binary choice models that is applicable, here). We assume that a country chooses PTA membership only if it gains in welfare and, accordingly, a PTA will be formed only if all members gain (see also BB). Similarly, accession of a country to an existing PTA will only take place if both the incumbent(s) and the entrant(s) expect to be better off with a PTA enlargement. Formally, we can write $\mathbf{PTA}^* = \min(\Delta U_1, \Delta U_2, \dots, \Delta U_m)$ with ΔU denoting the membership-to-non-membership utility differential of the 1, 2, ..., m (potential) members of a PTA. The observable variable \mathbf{PTA}_{ij} takes the value 1 if two countries are members of the same PTA (indicating $\mathbf{PTA}_{ij}^* > 0$), and 0 otherwise (indicating $\mathbf{PTA}_{ij}^* \leq 0$). In vector form (vectors and matrices are in bold), the unobservable utility differential is determined by the following process

$$\begin{aligned}\mathbf{PTA}^* &= \mathbf{X}\boldsymbol{\beta} + \boldsymbol{\varepsilon} \\ \mathbf{PTA} &= \mathbf{1}[\mathbf{PTA}^* > \mathbf{0}],\end{aligned}\tag{1}$$

where \mathbf{PTA} , \mathbf{PTA}^* , $\mathbf{1}$, $\mathbf{0}$, and $\boldsymbol{\varepsilon}$, are $n \times 1$ vectors with n denoting the number of country-pairs. \mathbf{X} is the $n \times k$ matrix of explanatory variables including the constant and $\boldsymbol{\beta}$ is a $k \times 1$ vector of unknown parameters.

In principle, one could estimate the model in (1) by a linear probability model, where the binary variable \mathbf{PTA} is regressed on the explanatory variables determining \mathbf{PTA}^* . However, there are well-known problems associated with this approach. Among those, the most important ones are (i) that the error term is then necessarily heteroskedastic which leads to inefficient test statistics and (ii) that the predicted probabilities of PTA membership can be smaller than zero or larger than unity (see Greene, 2003). Existing research on the determinants of PTA membership avoids these problems by deploying non-linear probability models based on the assumption of normally (or log-normally) distributed disturbances.

Magee (2003) and BB estimate probit models, where ε_{ij} is identically and independently distributed following the normal distribution $N(0, \sigma_\varepsilon^2)$. However, these models assume that PTA memberships are independent of each other. But the latter is at odds

with Baldwin’s domino theory and also with the models in Frankel, Stein, and Wei (1995, 1998) and BB. If PTA memberships are interdependent, we cannot obtain consistent estimates of β from estimating (1). Then, the model to be estimated in vector form reads

$$\begin{aligned} \mathbf{PTA}^* &= \rho \mathbf{W} \cdot \mathbf{PTA}^* + \mathbf{X}\beta + \varepsilon \\ \mathbf{PTA} &= \mathbf{1}[\mathbf{PTA}^* > \mathbf{0}], \end{aligned} \tag{2}$$

where ρ is an unknown parameter and \mathbf{W} is an $n \times n$ matrix of known entries that determines the form of the interdependence across country-pairs. Since Hypothesis 2 suggests that the interdependence declines in trade costs, the entries of \mathbf{W} should inversely depend on the trade costs between the country-pairs. Hence, interdependence is captured by a separate explanatory variable, reflecting an inverse-trade cost weighted average of the dependent variable. In econometric terms, the latter is referred to as a spatial lag.

Unfortunately, there are two serious problems in limited dependent variable models with a spatial lag. First, such a data generating process leads to multiple integrals in the likelihood function. Second, the error term is likely heteroskedastic rendering the associated parameter estimates inconsistent if this is not accounted for (see McMillen, 1992). Hence, the spatial binary choice model for interdependent PTA memberships cannot be estimated simply by maximum likelihood as binary choice models usually are.

3.2 Cures: binary dependent variable models with a spatial lag

As indicated before, a model of cross-sectional dependence with an endogenous spatial lag is suited for our problem, since interdependence in the PTA membership-induced welfare effects monotonously declines in trade costs within the empirically relevant range. Trade costs are known to increase in geographical distance.⁷ A spatially lagged dependent variable is the spatial equivalent to a time-lagged dependent variable. There is a large body of research on the estimation of models with a spatial lag of a continuous dependent

⁷See Anderson and van Wincoop (2003) or Baier and Bergstrand (2005), for recent applications of gravity models, where trade costs are associated with distance and common borders.

variable using either maximum likelihood (Anselin, 1988) or generalized method of moments (Kelejian and Prucha, 1999). However, much less research has been undertaken to estimate models with binary dependent variables.

McMillen (1992) is credited with being one of the first to provide an easily tractable solution to the problem. He proposes an EM algorithm which replaces the binary dependent variable with the expectation of the underlying continuous latent variable. This variable is then treated as a standard continuous one in the maximum likelihood estimation. The procedure is repeated until convergence. However, several problems arise with McMillen's model (LeSage, 1997, 2000). First, the method prohibits the use of the information matrix approach to determining the precision of the parameter estimates. In particular, the framework rules out estimates of dispersion for the parameter of the spatial lag, which is central to our analysis. Also, the confidence bounds around the other parameters are typically too small. Second, it is not suited for large-scale problems such as ours, covering more than 15,000 cross-sectional observations. Third, it requires knowledge about the functional form or variables involved in the non-constant variance relationship. Case (1992) derives an alternative estimator to McMillen's. But hers is only applicable to data-sets where the observations can be grouped into regions whose errors are strictly independent of each other (LeSage, 2000).

These problems can be overcome by relying on the Markov chain Monte Carlo method as proposed by LeSage (1997, 2000). This approach is also referred to as Gibbs sampling. The principal advantages are its suitability for large-scale problems of spatial dependence such as ours and its flexibility regarding the possible underlying heteroskedasticity of the error term. It specifies the complete conditional distributions for all parameters in the model. Sampling from these distributions then obtains a large set (a chain) of parameter draws. The corresponding estimates can be shown to converge in the limit to the joint posterior distribution of the parameters (Gelfand and Smith, 1990, LeSage, 2000).

Formally, the empirical model is a Bayesian heteroskedastic spatial autoregressive

probit model that can be written as follows:

$$\begin{aligned}\mathbf{PTA}^* &= \rho \mathbf{W} \cdot \mathbf{PTA}^* + \mathbf{X}\boldsymbol{\beta} + \boldsymbol{\varepsilon} \\ \boldsymbol{\varepsilon} &\sim N(0, \sigma^2 \mathbf{V}), \\ \mathbf{V} &= \text{diag}(v_1, v_2, \dots, v_n),\end{aligned}\tag{3}$$

To allow for heteroskedasticity, the elements of $\boldsymbol{\varepsilon}$ exhibit a non-constant variance, where $\sigma^2 v_i$ denotes the variance for observation i .

In a Bayesian approach, one applies Bayes' rule to learn about the unknown parameters based on the data. In such a framework, the *posterior density* of the parameters (and hence the parameters that fit the data best) is determined by the product of the likelihood function and the *prior density*. The latter two hinge upon assumptions. In our application, the likelihood function reads

$$L(\rho, \boldsymbol{\beta}, \sigma^2, \mathbf{V}, \mathbf{y}, \mathbf{W}) = \sigma^{-n} \prod_{i=1}^n (1 - \rho \mu_i) \prod_{i=1}^n v_i^{-1/2} \exp \left[- \sum_{i=1}^n \frac{\varepsilon_i^2}{2\sigma^2 v_i} \right],\tag{4}$$

where ε_i is the i th element of $(\mathbf{I}_n - \rho \mathbf{W})\mathbf{y} - \mathbf{X}\boldsymbol{\beta}$. The determinant $|\mathbf{I}_n - \rho \mathbf{W}|$ is written as $\prod_{i=1}^n (1 - \rho \mu_i)$, with μ_i denoting the eigenvalues of the matrix \mathbf{W} . Priors have to be formed about the set of parameters to be estimated: ρ , $\boldsymbol{\beta}$, σ^2 , and (v_1, v_2, \dots, v_n) . The latter relative variance terms are assumed to be fixed but unknown parameters. Estimating these n additional parameters seems to be problematic regarding the loss of degrees of freedom. However, the Bayesian approach relies on informative priors about the parameters v_i . In particular, an independent $\chi^2(r)/r$ distribution is assumed about the priors on (v_1, v_2, \dots, v_n) . The χ^2 distribution relies on a single parameter, r . Hence, the n parameters v_i in the model can be estimated by relying on a single parameter r in the estimation.⁸ The priors on $\boldsymbol{\beta}$ are assumed to be normally distributed with mean zero and variance 10^{12} (hence, these priors are relatively uninformative), the prior on σ^2 is

⁸Lindley (1971) used this type of prior for cell variances in an analysis of variance problem, and Geweke (1993) in modeling heteroscedasticity and outliers in the context of linear regression. Our runs for the heteroskedastic models rely on $r = 4$.

proportional to $1/\sigma$, and the priors on ρ and r are assumed to be constant. It is assumed that all priors are independent of each other.

The posterior density kernel to this model is given by the product of the likelihood function and the priors as assumed above. Unfortunately, this leads to an analytically intractable joint distribution. However, the conditional distributions for the parameters of interest can be set forth (see Albert and Chib, 1993, and Geweke, 1993, for the foundations). LeSage (1997, 2000) derives the conditional distributions for discrete choice models with spatial dependence as ours (see Appendix E for details). These conditional distributions can be used to compute posterior moments for all functions of interest using Gibbs sampling. Therein, we rely on 10,500 draws. Below, we will estimate the first and second moments of the distribution of these draws for parameter inference. These moments are computed after skipping 500 burn-ins. Hence, 500 draws are dropped to ensure that there is no systematic information left in the random numbers generation process for the remaining 10,000 draws. If there is a high autocorrelation in the Monte Carlo chain for each parameter, proper inference on the standard deviation may require dropping further draws from the chain (see Raftery and Lewis, 1992a,b, 1995). We use the Monte Carlo chain estimates then also to compute the first and second moments of the marginal effects to compare the outcome of the spatial probit model of PTA formation to its simple probit counterpart.

4 Empirical analysis for 15,753 country-pairs

4.1 Specification

In the empirical analysis, we rely on a specification that is similar to the one in BB. We use the following variables (the expected signs are in parentheses):

- NATURAL (+) measures the log of the inverse of the great circle distance between two trade partners' capitals.
- DCONT(+) is a dummy variable that takes the value one if two countries are located

at the same continent and zero else.⁹

- REMOTE = DCONT · 0.5{log[$\sum_{k \neq j} \text{Distance}_{ik}/(N-1)$] + log[$\sum_{k \neq i} \text{Distance}_{kj}/(N-1)$]} (+) is a country pair's remoteness from the rest of the world.
- total bilateral market size RGDPsum = log(RGDP_{*i*} + RGDP_{*j*}) (+) with RGDP_{*i*}, RGDP_{*j*} denoting the real GDP of countries *i, j*.
- RGDPsim = log{1 - [RGDP_{*i*}/(RGDP_{*i*} + RGDP_{*j*})]² - [RGDP_{*j*}/(RGDP_{*i*} + RGDP_{*j*})]²} (+).¹⁰
- DKL = |log(RGDP_{*i*}/POP_{*i*}) - log(RGDP_{*j*}/POP_{*j*})| (+) is the absolute difference in real GDP per capita.¹¹
- SQDKL = DKL² (-).
- DROWKL = 0.5{|log($\sum_{k \neq i} \text{RGDP}_k / \sum_{k \neq i} \text{POP}_k$) - log(RGDP_{*i*}/POP_{*i*})| + |log($\sum_{k \neq j} \text{RGDP}_k / \sum_{k \neq j} \text{POP}_k$) - log(RGDP_{*j*}/POP_{*j*})|} (-) is the relative factor endowment difference between the rest of the world and a given country-pair. We set up the database such that every country-pair arises only once. With 178 countries in the sample, there are then $178 \cdot (178 - 1) / 2$ unique pairs in the sample.¹²

Table 6 in Appendix B provides a summary of the descriptive statistics of the dependent and the explanatory variables in use.

However, our primary interest is on interdependence. Hence, we include the variable **W · PTA** in our model as outlined before. For this, we need to specify the weighting

⁹BB use only NATURAL instead of NATURAL and DCONT together. However, our results indicate that both of them should be included.

¹⁰BB use the absolute value of the difference in log real GDP of two countries instead. Consequently, the expected sign for their parameter is negative.

¹¹Already Kaldor (1963) pointed to the high correlation of capital-labor ratios and real GDP per capita. Capital stock data for a large country sample as ours are not available. Even perpetual inventory method based estimates thereof as in BB can not be derived due to missing data on gross fixed capital formation and investment deflators (see Leamer, 1984). If interdependence matters, the enormous loss of observations due to the use of capital stock values can not be justified. With a serious decline in observations, the problem of interdependence could not be consistently accounted for anymore, leading to eventually biased probit estimates.

¹²Hence, US-Canada and Canada-US are treated as being the same pair.

matrix \mathbf{W} whose elements should be inversely related to the distance (trade costs) between country-pairs ℓ and m . Suppose that country-pair ℓ consists of economies i and j and country-pair m of countries h and k . We define the distance between pairs ℓ and m as $\text{Distance}_{\ell m} = (\sum_{\iota} \sum_{\kappa} \text{Distance}_{\iota\kappa}) / 4$ with $\iota = i, j$ and $\kappa = h, k$. We employ three alternative weighting schemes. All of them exhibit elements $\omega_{\ell m}$ that are based on $w_{\ell m} = e^{-\text{Distance}_{\ell m}/500}$ if $\text{Distance}_{\ell m} < 2000$. We use a cut-off distance of 2000 kilometers to avoid problems associated with an excessive memory requirement for matrix elements that are close to zero anyway.¹³ We divide the exponent in $w_{\ell m}$ to ensure that the decay of the interdependence is slow enough. We use alternative weights in the sensitivity analysis later on. In general, \mathbf{W} is row-normalized for econometric reasons such that $\omega_{\ell m} = w_{\ell m} / \sum_m w_{\ell m}$. Then, we know for the parameter measuring the strength of interdependence that $0 \leq |\rho| \leq 1$ is required for proper inference. An alternative weighting scheme sets all elements $\omega_{\ell m} = 0$ if there is no overlap in the two country-pairs ℓ and m . This captures interdependence within PTAs so that the corresponding spatial lag parameter reflects the incentive to join other countries in a PTA. There, we focus on the interdependence in membership for country-pairs such as Canada-US and US-Mexico. Let us denote this weighting matrix as $\mathbf{W}_{\text{direct}}$. The third considered matrix $\mathbf{W}_{\text{indirect}} = \mathbf{W} - \mathbf{W}_{\text{direct}}$ captures the interdependence of non-overlapping country-pairs such as US-Mexico and Germany-France. In the theoretical section, we illustrated that there should be a stronger incentive for joining other trading partners in a PTA (interdependence within PTAs) than creating one in response to others (interdependence across PTAs). The distinction between $\mathbf{W}_{\text{direct}}$ and $\mathbf{W}_{\text{indirect}}$ and their alternative use enables an empirical assessment of this hypothesis. According to Hypothesis 3 derived from the BB model, we would expect $\rho, \rho_{\text{direct}}, \rho_{\text{indirect}} > 0$, hence, interdependence matters and trade diversion through foreign PTA formation stimulates a country's probability to join/form a PTA with others. Furthermore, we hypothesize that $\rho_{\text{indirect}} < \rho_{\text{direct}}$, hence, the incentive to join other countries in their PTA is stronger than forming one with other outsiders in

¹³Note that it is impossible to handle (invert, transpose, and even store) a full $15,753 \times 15,753$ for any modern personal computer.

response to foreign PTA formation.

We have put great effort into the efficiency of the implementation following LeSage (1999a,b). Just to portray the size of the problem: the sheer construction of the matrix \mathbf{W} by using a standard loop (running over $15,753 \times 15,753$ country-pairs) in MATLAB takes about 48 hours.¹⁴ The estimation of the spatial probit model with heteroskedasticity-robust standard errors based on \mathbf{W} and 10,500 Monte Carlo draws takes more than 60 hours.

4.2 Estimation results

We estimate a standard probit model similar to BB and three different spatial probit models based on the alternative weighting matrices \mathbf{W} , $\mathbf{W}_{\text{indirect}}$, and $\mathbf{W}_{\text{direct}}$. In the spatial models, we allow for heteroskedastic disturbances (see Section 5 for the results of the corresponding homoskedastic models). Table 1 summarizes our findings.

— Table 1 —

The standard probit model obtains results that are similar to the ones in BB. Countries that are closer to each other in geographical terms and that are located at the same continent exhibit a higher probability of PTA membership ($\hat{\beta}_{\text{NATURAL}} > 0$, $\hat{\beta}_{\text{DCONT}} > 0$). Country-pairs that are relatively remote from the rest of the world will more likely form a PTA ($\hat{\beta}_{\text{REMOTE}} > 0$). Also larger and more similarly sized economies tend to form a PTA more likely than others ($\hat{\beta}_{\text{RGDPsum}} > 0$, $\hat{\beta}_{\text{RGDPsim}} > 0$). Countries with dissimilar relative factor endowments are more likely inclined towards forming a PTA than similar ones ($\hat{\beta}_{\text{DKL}} > 0$), and the squared relative factor endowment difference variable enters negatively ($\hat{\beta}_{\text{SQDKL}} < 0$). These point estimates are in line with those of BB. However, we do not find a significantly negative effect of the difference in relative factor endowments from the rest of the world. Rather, the corresponding point estimate is positive ($\hat{\beta}_{\text{DROWKL}} > 0$) but not significantly different from zero.¹⁵ The pseudo- R^2 of this

¹⁴The hardware in use is a Fujitsu Siemens PC with 2 gigabyte RAM and a 3.2 gigahertz processor.

¹⁵Notice that BB do not include SQDKL and DROWKL simultaneously.

model amounts to 0.229 which is smaller than the one estimated in the much narrower country sample of BB.

Let us now turn to the spatial models that account for interdependence in PTA membership. Consider first the benchmark spatial model based on \mathbf{W} in the second column labeled 'Spatial, \mathbf{W} ' of Table 1. Interestingly, we find that, with a few exceptions (DKL and DROWKL), there is only a minor change in the parameters of the economic fundamentals used in the simple probit. But we identify a significant, large, positive effect of the interdependence parameter $\hat{\rho} = 0.805$. The finding of $\hat{\rho} > 0$ and its significance supports Hypotheses 1 and 2 at the same time. As we will see below, this has an important consequence for the marginal effects.

The significance of the spatial interdependence also leads to a higher value of the pseudo-log-likelihood evaluated at the estimated parameters than in the simple probit model (see LeSage, 1997, 2000; a usual log-likelihood value is not available for the spatial models). The Raftery and Lewis (1992a,b, 1995) diagnostics based on 10,000 draws (after dropping the burn-ins) suggested dropping every second draw for all models estimated to avoid an excessive autocorrelation in the Monte Carlo chain.¹⁶ Accordingly, the first and second moments of the posterior parameter distributions reported in the table are based on 5,000 draws only. The reported diagnostics indicate that there are enough draws and burn-ins for proper inference. Note that the ratio between the total number of draws needed to achieve an accuracy for testing at 5 percent and the ones required under identically and independently distributed (i.i.d.) draws exhibits a value that is much lower than 5, as required for proper convergence. Also, a set of further convergence diagnostics suggested by Geweke (1992) leads to this conclusion but is suppressed here in order to save space.

To shed light on the relative strength of interdependence for joining a PTA versus forming one with other outsiders, we run alternative regressions using $\mathbf{W}_{\text{indirect}}$ and $\mathbf{W}_{\text{direct}}$. Indeed, we find that $\hat{\rho}_{\text{direct}} > \hat{\rho}_{\text{indirect}}$, which lends support to Hypothesis 3.

¹⁶An excessive autocorrelation in the chain unnecessarily inflates the standard deviation of the parameters.

While the spatial probit model parameter estimates per se indicate that interdependence matters, they do not allow for drawing conclusions about their difference to the simple counterpart in quantitative terms. For this, we need to consider the predicted probabilities of membership and the marginal effects of the variables in the model. Table 2 summarizes descriptive statistics of the estimated probabilities of membership for the different models.

-- Table 2 --

In our sample, the average country-pair exhibits a probability of PTA membership of about 14 percent. This seems reasonable since about 14 percent of the country-pairs are in fact PTA members. However, the standard deviation of the estimated probability is about as large as the mean, and the minimum and maximum values amount to 0 and about 97 percent, respectively. Among the spatial models, the one based on the weighting matrix \mathbf{W} is best comparable to the simple probit model. Note that the elements in \mathbf{W} are set to zero for PTA pairs that are more than 2000 kilometers away from each other. Hence, a zero impact is assumed for such pairs on each other. The average predicted probability for this spatial model amounts to about 11 percent which is somewhat smaller than that for the simple probit. The standard deviation and the minimum are similar to the simple probit model, and the maximum amounts to about 100 rather than 97 percent. The other models only rely on a sub-matrix of \mathbf{W} . However, $\mathbf{W}_{\text{direct}}$ contains only about 0.03 percent of the non-zero entries of \mathbf{W} whereas $\mathbf{W}_{\text{indirect}}$ contains the remaining 99.97 percent. The average of the \mathbf{W} -based estimated membership probabilities comes close to the simple probit model, and the corresponding average based on $\mathbf{W}_{\text{indirect}}$ is closer to the simple probit than the one based on $\mathbf{W}_{\text{direct}}$ (see Table 2).

The mean in Table 2 is informative about the prediction differences between the spatial and the simple probit models for a randomly drawn country-pair in the sample. However, this difference can be quite large for a specific pair. In our sample, the average difference between the simple probit model and the spatial model based on \mathbf{W} amounts to about 4 percentage points. The corresponding standard deviation is 7 percentage points, and

the minimum and maximum differences amount to about -37 and 41 percentage points, respectively. Hence, for one country-pair we expect a probability of being a PTA member that is 38 percentage points lower than in the simple probit. For another one, the probability of forming a PTA is 41 percentage points higher than in the simple probit model. Table 3 provides details on the extreme spatial-to-simple probit model differences when using the spatial model based on \mathbf{W} .

-- Table 3 --

In the top panel of the table, the probability of PTA membership predicted by the simple probit model is quite high, and it is much lower in the spatial model. Obviously, the largest negative deviations of the spatial model from the simple one arise for Djibouti-Somalia (-37 percentage points), Oman-Saudi Arabia (-35 percentage points), India-Iran (-34 percentage points), Iran-Saudi Arabia (-33 percentage points), and Israel-Saudi Arabia (-33 percentage points). By and large, these countries are located at or close to the Arabian Peninsula. The reason for this finding is that the parameters in the spatial probit model are somewhat lower than in its simple counterpart. This leads to smaller predicted membership probabilities. Since there are only a few PTAs in the neighborhood of these countries, the lower direct effect on PTA membership probabilities of these economies is compensated only to a minor extent by the effects of interdependence.

In the bottom panel of the table, the opposite holds true. There, the predictions of the simple probit model tend to be low (except for Belize-Nicaragua, where the predicted membership probability is higher than 50 percent) whereas those of the spatial model are much higher. The highest positive deviations of the predicted membership probabilities of the spatial probit from its simple counterpart arise for Aruba-Haiti (41 percentage points), Bahamas-Haiti (37 percentage points), Haiti-Netherlands Antilles (35 percentage points), Belize-Nicaragua (34 percentage points), and Haiti-Nicaragua (31 percentage points). Notice that these countries belong to the Caribbean with numerous PTAs in the neighborhood. Accordingly, the lower direct effect on PTA membership probabilities is more than compensated by the effects of interdependence for these economies.

Furthermore, let us consider the marginal effects of the variables in the model. We compute these effects for bilateral economic size and the geographical variables, primarily for the sake of comparison among the non-spatial and the spatial models. Table 4 summarizes the corresponding findings. The figures given there are marginal effects associated with a one percent increase in the respective variables. The findings suggest the following conclusions. First, the positive marginal effect for NATURAL is very similar across the board. Second, the remaining marginal effect estimates tend to be considerably lower with the spatial models than with the non-spatial one. For instance, the estimate of REMOTE in the \mathbf{W} -based model amounts to only 41 percent of its non-spatial counterpart.

Hence, we may conclude that accounting for interdependence of PTAs matters for both the predicted probabilities of forming a PTA and the estimates of the marginal effects of the economic fundamentals.

5 Sensitivity analysis

In the sequel, we provide insights into the robustness of our findings in various ways as far as this is computationally feasible. Since the paper’s focus is on interdependence, we confine ourselves to a summary of the corresponding estimates of the parameters ρ , ρ_{indirect} , and ρ_{direct} . We discuss the qualitative and quantitative changes in these parameters in the following and report the corresponding means and standard deviations of the Markov chain as well as the log-likelihood values in Table 5 below.

(i) *Altering the decay of the spatial weights matrix:* In the previous analysis, the elements of the spatial weights matrices were based on $w_{\ell m} = e^{-\text{Distance}_{\ell m}/\psi}$ with $\psi = 500$. The corresponding distance-related decay of interdependence was such that $w_{\ell m}$ was 0.67, 0.37, 0.14, and 0.02 for distances among country-pairs of 200, 500, 1000, and 2000 kilometers, respectively. In an alternative set of results, we have used alternative decay parameters of $\psi = 1000$ and $\psi = 250$. For instance, with $\psi = 1000$ the entries $w_{\ell m}$ become 0.82, 0.61, 0.37, and 0.14 for distances of 200, 500, 1000, and 2000 kilometers, respectively.

In contrast, with $\psi = 250$ and the same distance examples the entries $w_{\ell m}$ become 0.45, 0.14, 0.02, and 0.0003, respectively. With the weights matrices based on $\psi = 1000$, we find that $\hat{\rho} = 0.903$ and $\hat{\rho}_{\text{indirect}} = 0.771 < \hat{\rho}_{\text{direct}} = 0.961$. For $\psi = 250$, $\hat{\rho} = 0.844$ and $\hat{\rho}_{\text{indirect}} = 0.690 < \hat{\rho}_{\text{direct}} = 0.950$. Hence, the change in the decay parameter ψ affects the parameter estimates reflecting interdependence in quantitative terms. However, the qualitative results are unchanged ($\hat{\rho} > 0$ and $\hat{\rho}_{\text{indirect}} < \hat{\rho}_{\text{direct}}$).

— Table 5 —

(ii) *Altering the cut-off level in the spatial weights matrix:* As indicated before, we have so far used a cut-off level of 2000 kilometers. Hence, all PTAs with a distance lower than that were allowed to have an impact on a country-pair's probability to preferentially eliminate tariffs.¹⁷ For sheer memory reasons, it is not possible to increase that cut-off point to, say, 3000 kilometers. However, we ran regressions based on a lower cut-off value of 1000 kilometers (using the original decay parameter value of $\psi = 500$) to investigate the qualitative robustness of our findings in this regard. Since a lower cut-off value increases the sparseness of the spatial weights matrix, the computing time for model estimation is reduced as well. We find that $\hat{\rho} = 0.903$ and $\hat{\rho}_{\text{indirect}} = 0.770 < \hat{\rho}_{\text{direct}} = 0.961$. Hence, the findings respond quantitatively to the choice of an alternative distance cut-off value but the qualitative results are again unchanged.

(iii) *Altering the skewness of the χ^2 -distribution of the v_i :* We have used a skewness parameter $r = 4$ for the χ^2 distribution of the n parameters v_i , determining the degree of heteroskedasticity in the model. The parameters of non-linear probability models can react quite sensitively to the heteroskedasticity of the disturbances. We assess the robustness of our findings with respect to alternative choices for r , namely $r = 2$ and $r = 20$. Furthermore, we estimate a model that assumes homoskedastic disturbances such that $v_i = 1$ for all $i = 1, \dots, n$. When using $r = 2$ instead of $r = 4$, $\hat{\rho} = 0.805$ and $\hat{\rho}_{\text{indirect}} = 0.635 < \hat{\rho}_{\text{direct}} = 0.975$. With $r = 20$ we obtain estimates $\hat{\rho} = 0.812$

¹⁷Note that the cut-off is based on two countries' average (not the maximum) distance to two economies. This average could be less than 2000 kilometers even though the maximum distance of one economy to another entering that average could be much more than that.

and $\hat{\rho}_{\text{indirect}} = 0.632 < \hat{\rho}_{\text{direct}} = 0.978$. Assuming that the residuals are homoskedastic leads to $\hat{\rho} = 0.748$ and $\hat{\rho}_{\text{indirect}} = 0.571 < \hat{\rho}_{\text{direct}} = 0.972$. These findings suggest that heteroskedasticity is indeed present in the sample and should not be omitted. However, in qualitative terms, our original conclusions would not be changed with a homoskedastic model. And they are robust to the choice of higher and lower skewness parameters of the χ^2 -distribution about v_i .

(iv) *Augmenting the specification by political variables:* Both Magee (2003) and BB put forward evidence that political variables could explain some of the variation in PTAs. We use a set of potentially relevant political variables made available through the Polity IV project (Marshall and Jaggers, 2002). Among a few others, this data-set forms a cornerstone for empirical research in political science. In particular, we employ the following variables in our augmented specification: a numeric democracy score (defined on a range from 0 to 10, where 10 reflects a high degree of general openness of political institutions); a numeric autocracy score (defined on a range from 0 to 10, where 10 reflects a high degree of closedness of political institutions);¹⁸ and a numeric score of political competition (a high score reflects a high development of institutional structures for political expression and a high accessibility of these structures by non-elites). We use both the minimum value and the maximum value of each variable at the country-pair level separately. This is as informative as using averages and/or absolute differences in these variables. A positive sign of the respective minimum value and a negative one of the maximum value at the country-pair level indicates that similarity in a particular political score is associated with a higher probability of participating in a PTA. According to previous research, we would expect more democratic, and politically open societies to be more favorable of trade liberalization than autocratic, closed ones. It turns out that, indeed, the difference between the minimum and maximum coefficient of political variables is significantly negative. Hence, a greater similarity of political systems increases the probability to engage in a PTA. The political variables enter significantly in our models. For example, four (five) of the six co-

¹⁸Note that the democracy score is not just the inverse of the autocracy score. While the two scores are negatively correlated as expected, the partial correlation coefficient amounts to only -0.464 . Hence, there is enough variation in the data to include them simultaneously.

efficients are significant at five (ten) percent in a simple probit model. While we suppress a detailed summary of the results for all coefficients,¹⁹ we report our findings regarding the interdependence parameters of interest. With the augmented specification, we estimate $\hat{\rho} = 0.888$ and $\hat{\rho}_{\text{indirect}} = 0.701 < \hat{\rho}_{\text{direct}} = 0.986$. Accordingly, we may conclude that political variables play a role in determining a country-pair's probability of engaging in a PTA, but their omission does not affect the qualitative result that interdependence matters and is more important for joining other countries in a PTA (interdependence within PTAs) rather than forming PTAs in response to others abroad (interdependence across PTAs).²⁰

(v) *Treating only customs unions and free trade areas but not other preferential trade arrangements as PTAs:* Of the covered PTAs notified to the WTO there are 17 PTAs which are neither customs unions nor free trade areas.²¹ Since they represent a much weaker liberalization of trade barriers, they might affect the point estimates of the interdependence in PTA memberships. We assess the robustness of our findings by ignoring the mentioned 17 PTAs (i.e., setting the PTA dummy at one only for customs unions and free trade areas). Obviously, this does not affect our sample size. We find point estimates of our interdependence parameters of interest of $\hat{\rho} = 0.763$ and $\hat{\rho}_{\text{indirect}} = 0.604 < \hat{\rho}_{\text{direct}} = 0.959$, which support our original results.

(vi) *Using the same sample of country-pairs as Baier and Bergstrand (2004):* In a final step, we focus on the same subset of 1,431 country-pairs as BB to see whether there is strong interdependence among these pairs as well. However, we have to say upfront that the insights from this experiment are limited. With interdependence in the world economy in general, a restriction to such a small subset of country-pairs may lead to a severe bias in the interdependence parameters. Hence, the point estimates should be interpreted with caution. Yet, with the estimates of $\hat{\rho} = 0.774$ and $\hat{\rho}_{\text{indirect}} = 0.552 < \hat{\rho}_{\text{direct}} = 0.847$ we

¹⁹They are available from the authors upon request.

²⁰Note that the augmentation of the specification by political variables leads to a loss of 4,727 observations. Hence, we should be careful with a quantitative comparison of the interdependence parameters with the ones in the baseline models summarized in Table 1.

²¹These are AFTA, Bangkok Agreement, CAN, CEMAC, COMESA, EAC, ECO, GCC, GSTP, LAIA, MSG, Laos and Thailand, PTN, SAPTA, SPARTEKA, TRIPARTITE, UEMONA.

still find a similar qualitative pattern as before.

Accordingly, we may conclude that there is a robust indication of a positive interdependence in PTA membership in the world economy. The formation of PTAs generates a particularly strong incentive for non-distant outsiders to join, and there is a somewhat lower but still positive incentive for outsiders to form their own PTA.

6 Conclusions

This paper puts forward novel empirical insights about the determinants of preferential trade agreement (PTA) memberships. The focus is on the interdependence of PTA memberships in the world economy. We derive the following three testable hypotheses regarding interdependence: (i) the formation of foreign PTAs generates an incentive to lower tariffs preferentially for a country-pair to reduce the welfare loss from trade diversion; (ii) this incentive declines in the distance to foreign PTAs since the associated trade diversion is then lower; (iii) the incentive is stronger for joining other countries in a PTA (interdependence within PTAs) than it is for forming a PTA with other outsiders (interdependence across PTAs).

These hypotheses are investigated in a huge sample, covering 15,753 country-pairs. We employ a Bayesian spatial discrete choice model that obtains consistent estimates with interdependent data. There is significant support for any of the hypotheses which seems to be very robust to the chosen sample, the set of explanatory variables and various model assumptions. We illustrate that interdependence does not only matter as such, but its ignorance seriously affects the estimates of the model parameters. We provide evidence that the estimated probabilities of PTA membership are biased upwards or downwards by up to 40 percentage points in our sample.

Appendix

A The theoretical model

The model underlying our theoretical analysis is equivalent to the one of Baier and Bergstrand (2004). Here, we describe the main features for convenience.

A.1 Consumers

The model consists of three continents and two countries on each of them. Each country i hosts a single representative consumer who derives utility from the consumption of goods (g) and services (s), based upon Cobb-Douglas preferences with constant income shares of γ and $(1 - \gamma)$, respectively. Each sector is characterized by a taste for diversity which is formally captured by Dixit and Stiglitz (1977) preferences with a constant elasticity of substitution. Let g_{iik} (s_{iik}) be the consumption in country i of differentiated good (service) k produced in the home country i ; $g_{ii'k}$ ($s_{ii'k}$) is consumption in country i of good (service) k produced in the foreign country on the same continent i' ; and g_{ijk} (s_{ijk}) is consumption in country i of good (service) k produced in each of the four foreign countries on other continents j . Furthermore, θ^g (θ^s) denotes the parameter determining the elasticity of substitution between varieties of goods (services), with $0 < \theta^g, \theta^s < 1$. Finally, let n_i^g (n_i^s) be the number of varieties of goods (services) produced in the home country, $n_{i'}^g$ ($n_{i'}^s$) the number of varieties of goods (services) produced in the foreign country on the same continent, and n_j^g (n_j^s) the number of varieties of goods (services) produced by a foreign country on another continent.

The utility function U_i is then given by:

$$U_i = \left[\sum_{k=1}^{n_i^g} g_{iik}^{\theta^g} + \sum_{k=1}^{n_{i'}^g} g_{ii'k}^{\theta^g} + \sum_{j \neq i, i'} \sum_{k=1}^{n_j^g} g_{ijk}^{\theta^g} \right]^{\frac{\gamma}{\theta^g}} \left[\sum_{k=1}^{n_i^s} s_{iik}^{\theta^s} + \sum_{k=1}^{n_{i'}^s} s_{ii'k}^{\theta^s} + \sum_{j \neq i, i'} \sum_{k=1}^{n_j^s} s_{ijk}^{\theta^s} \right]^{\frac{1-\gamma}{\theta^s}}, \quad (\text{A1})$$

Within a country, households and firms are assumed symmetric, which allows elim-

inating subscript k . There are two factors of production, labor L and capital K . Let w_i denote the wage rate of the representative worker in country i , and r_i the rental rate on capital per household. Goods trade is impeded by intra- (inter-)continental transport costs of the Samuelson-'iceberg'-type, where $a \in [0, 1]$ ($b \in [0, 1]$) represents the fraction of output exported by a country that is lost due to intra- (inter-)continental transport. Furthermore, there are ad valorem import tariff rates on goods and services, where $t_{ii'}$ and t_{ij} denote the tariff rates levied by country i .

The consumer maximizes Equation (A1) under the budget constraint

$$E_i = w_i L_i + r_i K_i + T_i, \quad (\text{A2})$$

where T_i is the tariff revenue which is redistributed to households in a lump sum fashion. This maximization yields a set of demand equations, which are omitted here for the sake of brevity.

A.2 Firms

All firms in the goods and services industry are assumed to produce under the same technology. Specifically, the output of goods (services) produced by a firm, denoted by g_i (s_i), requires k_i^g (k_i^s) units of capital and l_i^g (l_i^s) units of labor, as well as an amount of ϕ^g (ϕ^s) of fixed costs expressed in terms of the output of the good produced. The production functions are then given by:

$$g_i = (k_i^g)^{\alpha^g} (l_i^g)^{1-\alpha^g} - \phi^g \quad s_i = (k_i^s)^{\alpha^s} (l_i^s)^{1-\alpha^s} - \phi^s. \quad (\text{A3})$$

Firms maximize profits subject to the technology defined in Equation (A3), given the demand schedule derived in Section (A.1). In this model, profit maximization leads to a constant markup over marginal production costs and zero profits due to free entry and exit.

A.3 Parameter values

The parametrization of the theoretical model is identical to the one in Baier and Bergstrand (2004). The income share spent on goods is set to $\gamma = 0.5$. The parameters determining the elasticity of substitution between varieties of goods and among services are set to $\theta^g = \theta^s = 0.75$, implying an elasticity of substitution of $\epsilon^g = \epsilon^s = \frac{1}{1-\theta^g} = 4$.

Parameters of the production function are chosen as follows: $\alpha^g = \alpha^s = 0.5$, $\phi^g = \phi^s = 1$.

Factor endowments of labor and capital are chosen equal across all six countries, i.e., $L_i = K_i = 100$ for $i = 1, \dots, 6$.

All initial tariffs as well as all post-integration tariffs on nonmember imports are set equal to 0.3, i.e., $t_{ii'} = t_{ij} = 0.3$. A rationale for this choice is provided in Frankel (1997).

Intra- (inter-)continental transport costs are varied between zero and one with a grid size of 0.01, i.e., $a = [0, 0.01, 0.99]$ ($b = [0, 0.01, 0.99]$) as in Baier and Bergstrand (2004).

B Data sources

We use information on PTAs that are notified to the World Trade Organization. These data are augmented and corrected by using information from the CIA's World Fact Book and PTA secretariat homepages, and they are compiled to obtain a binary dummy variable reflecting the most recent available year, 2005. Explanatory variables are based on raw data from the World Bank's World Development Indicators. Specifically, we use real GDP figures at constant parent country exchange rates and population. Bilateral distances are based on the great circle distance between two countries' capitals (own calculations, using coordinates as available from the CIA World Fact Book). The following table summarizes the descriptive statistics of the dependent and independent variables employed in the empirical specification.

-- Table 6 --

Most importantly, about 14 percent of the 15,753 country-pairs in our data-set are members of the same PTA. About 21 percent of the pairs are intracontinental ones.

C Country coverage (178 economies)

Afghanistan, Albania, Algeria, Angola, Argentina, Armenia, Aruba, Australia, Austria, Azerbaijan, The Bahamas, Bahrain, Bangladesh, Barbados, Belarus, Belgium, Belize, Benin, Bermuda, Bolivia, Bosnia and Herzegovina, Botswana, Brazil, Brunei, Bulgaria, Burkina Faso, Burundi, Cambodia, Cameroon, Canada, Cape Verde, Central African Republic, Chad, Chile, China, Colombia, Comoros, Rep. Congo, Costa Rica, Cote d'Ivoire, Croatia, Cuba, Cyprus, Czech Republic, Denmark, Djibouti, Dominican Republic, Ecuador, Egypt, El Salvador, Equatorial Guinea, Eritrea, Estonia, Ethiopia, Fiji, Finland, France, French Polynesia, Gabon, The Gambia, Georgia, Germany, Ghana, Greece, Guatemala, Guinea, Guinea-Bissau, Guyana, Haiti, Honduras, Hong Kong (China), Hungary, Iceland, India, Indonesia, Islamic Rep. Iran, Iraq, Ireland, Israel, Italy, Jamaica, Japan, Jordan, Kazakhstan, Kenya, Rep. Korea, Kuwait, Kyrgyz Republic, Lao PDR, Latvia, Lebanon, Lesotho, Liberia, Libya, Lithuania, Luxembourg, Macao (China), FYR Macedonia, Madagascar, Malawi, Malaysia, Mali, Malta, Mauritania, Mauritius, Mexico, Fed. Sts. Micronesia, Moldova, Mongolia, Morocco, Mozambique, Myanmar, Namibia, Nepal, Netherlands, Netherlands Antilles, New Caledonia, New Zealand, Nicaragua, Niger, Nigeria, Norway, Oman, Pakistan, Palau, Panama, Papua New Guinea, Paraguay, Peru, Philippines, Poland, Portugal, Puerto Rico, Qatar, Romania, Russian Federation, Rwanda, Samoa, Sao Tome and Principe, Saudi Arabia, Senegal, Sierra Leone, Singapore, Slovak Republic, Slovenia, Somalia, South Africa, Spain, Sri Lanka, St. Lucia, Sudan, Suriname, Swaziland, Sweden, Switzerland, Syrian Arab Republic, Tajikistan, Tanzania, Thailand, Togo, Tonga, Trinidad and Tobago, Tunisia, Turkey, Turkmenistan, Uganda, Ukraine, United Arab Emirates, United Kingdom, United States, Uruguay, Uzbekistan, Vanuatu, RB Venezuela, Vietnam, Rep. Yemen, Zambia, Zimbabwe.

D PTA coverage (127 agreements)

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, Armenia and Ukraine, 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), Communauté Économique et Monétaire de l'Afrique Centrale (CEMAC), 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 Tunisia, 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,

Georgia and Ukraine, 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, Kyrgyz Republic and Ukraine, Kyrgyz Republic and Uzbekistan, Asociación Latinoamericana de Integración (ALADI, LAIA), Laos and Thailand, 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).

E Econometric issues

Following Albert and Chib (1993) and Geweke (1993), LeSage (1997, 2000) derives the conditional posterior distributions of the parameters of interest in the discrete choice

model with a spatial lag:

$$\begin{aligned}
p(\boldsymbol{\beta}|\rho, \sigma, \mathbf{V}) &= N[(\mathbf{X}'\mathbf{V}^{-1}\mathbf{X})^{-1}\mathbf{X}'\mathbf{V}^{-1}(\mathbf{I}_n - \rho\mathbf{W})\mathbf{y}, \sigma^2(\mathbf{X}'\mathbf{V}^{-1}\mathbf{X})^{-1}], \quad (\text{A4}) \\
p(\sigma|\boldsymbol{\beta}, \rho, \mathbf{V}) &\propto \sigma^{-(n+1)}e^{-\sum_{i=1}^n \varepsilon_i^2/(2\sigma^2 v_i)}, \\
p(\rho|\boldsymbol{\beta}, \sigma, \mathbf{V}) &\propto |\mathbf{I}_n - \rho\mathbf{W}|e^{-(1/2\sigma^2)(\boldsymbol{\varepsilon}'\mathbf{V}^{-1}\boldsymbol{\varepsilon})}, \\
p(v_i|\boldsymbol{\beta}, \rho, \sigma, \mathbf{V}_{-i}) &\propto (\varepsilon_i^2/\sigma^2 + r)/v_i,
\end{aligned}$$

where \propto indicates that the expression on the left-hand side is proportional up to a constant to the one on the right-hand side, and \mathbf{V}_{-i} indicates all elements except v_i .

The posterior distribution of \mathbf{PTA}^* conditional on the model parameters takes the form of a truncated normal distribution. The latter is derived by truncating the function $N[\widehat{\mathbf{PTA}}_i^*, \sum_j \omega_{ij}^2]$ from the right by zero if $\mathbf{PTA}_i = 0$ and from the left by zero if $\mathbf{PTA}_i = 1$. There, $\widehat{\mathbf{PTA}}_i^*$ is the predicted value of the i th row of \mathbf{PTA}_i^* , and $\sum_j \omega_{ij}^2$ denotes the variance of the prediction with ω_{ij} denoting the ij th element of $(\mathbf{I}_n - \rho\mathbf{W})^{-1}\boldsymbol{\varepsilon}$. The probability density function of the latent variable \mathbf{PTA}^* is:

$$f(\mathbf{PTA}_i^*|\rho, \boldsymbol{\beta}, v_i) \sim \begin{cases} N(\widehat{\mathbf{PTA}}_i^*, \sum_j \omega_{ij}^2), & \text{truncated at the left by 0 if } \mathbf{PTA}_i = 1, \\ N(\widehat{\mathbf{PTA}}_i^*, \sum_j \omega_{ij}^2), & \text{truncated at the right by 0 if } \mathbf{PTA}_i = 0. \end{cases} \quad (\text{A5})$$

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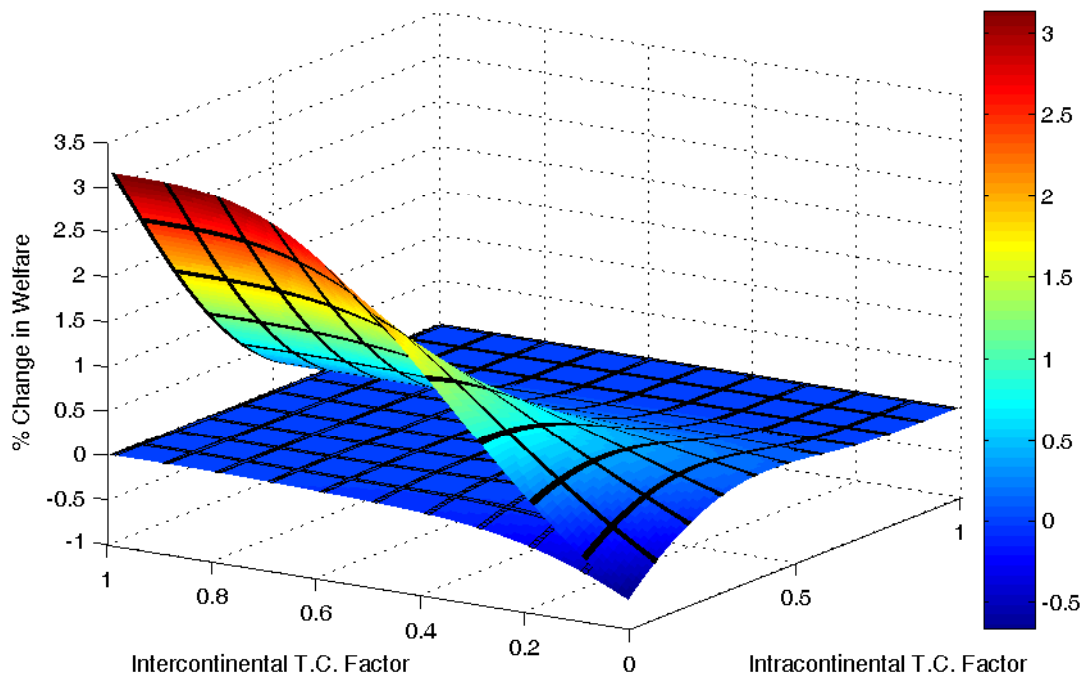


Figure 1: Net welfare gains from natural (top surface) and unnatural (bottom surface) PTAs. Natural PTAs among countries 1A and 2A, countries 1B and 2B, and countries 1C and 2C. Unnatural PTAs among countries 1A and 1B, countries 2A and 1C, and countries 2B and 2C.

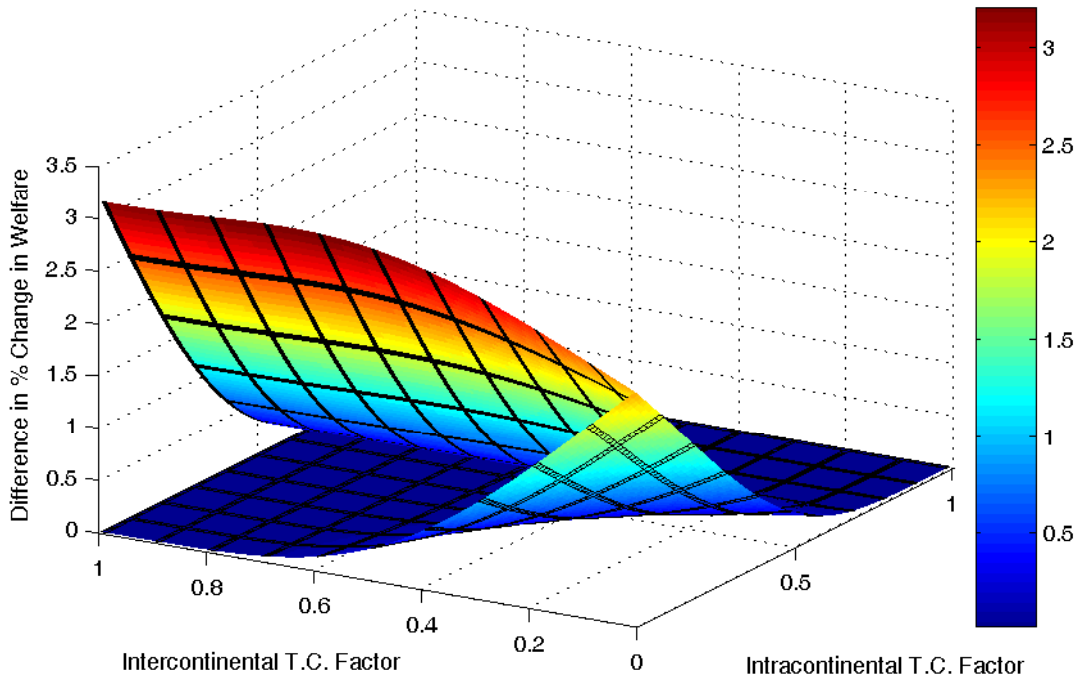


Figure 2: Top surface: Difference in net welfare gains from natural PTAs among countries 1A and 2A, countries 1B and 2B, and countries 1C and 2C, and those from PTAs only among countries 1B and 2B, and countries 1C and 2C. Bottom surface: Difference in net welfare gain from unnatural PTAs among countries 1A and 1B, countries 2A and 1C, and countries 2B and 2C, and those from PTAs only among countries 2A and 1C, and countries 2B and 2C.

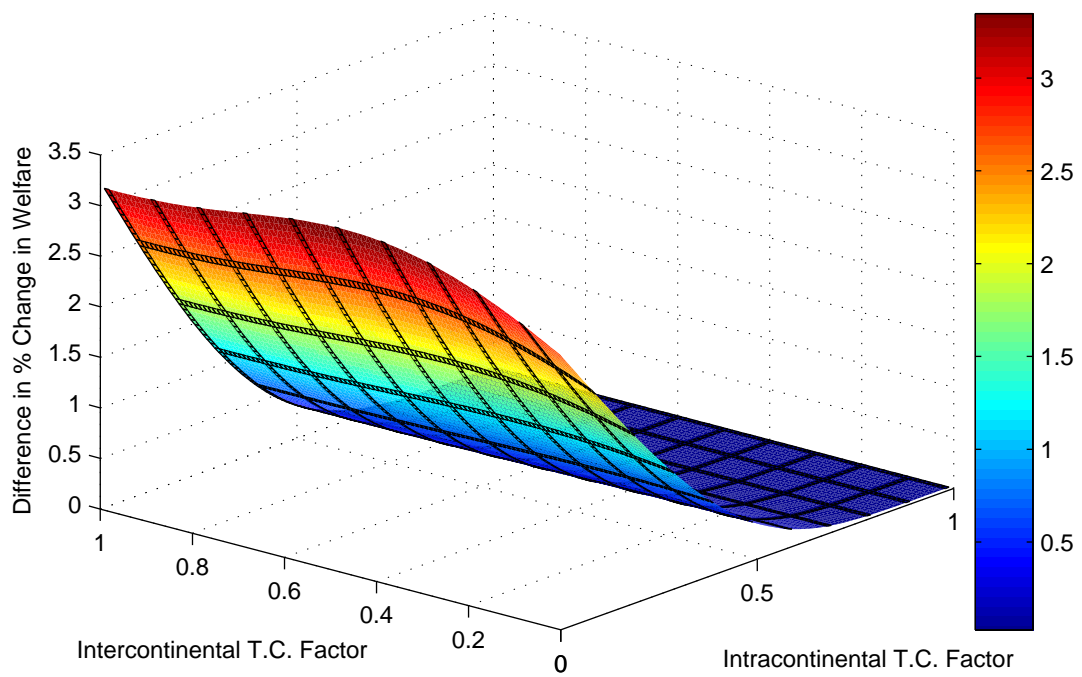


Figure 3: Difference in the net welfare gains from PTAs among countries 1A, 2A and 1C, and countries 2B and 2C, and those from unnatural PTAs among countries 1A and 1B, countries 2A and 1C, and countries 2B and 2C.

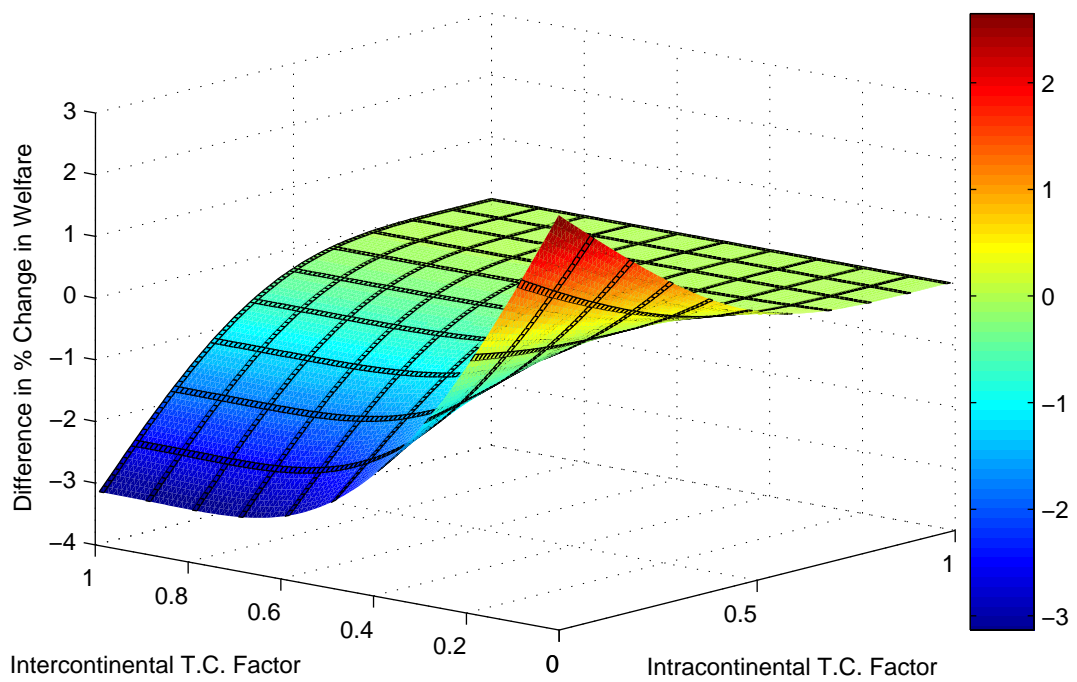


Figure 4: Difference in the net welfare gains from PTAs among countries $1A$, $1B$ and $2B$, and countries $1C$ and $2C$, minus those from natural PTAs among countries $1A$ and $2A$, countries $1B$ and $2B$, and countries $1C$ and $2C$.

Table 1: Probit results for the probability of a preferential trade agreement (non-spatial and spatial models)

Parameters	Theory	Non-spatial	Spatial, \mathbf{W}^c	Spatial, $\mathbf{W}_{\text{indirect}}^c$	Spatial, $\mathbf{W}_{\text{direct}}^c$
ρ	+		0.805*** 0.035	0.632*** 0.039	0.973*** 0.016
NATURAL	+	0.517*** 0.024	0.761*** 0.030	0.723*** 0.031	0.777*** 0.028
RGDPsum	+	0.191*** 0.009	0.128*** 0.011	0.139*** 0.011	0.111*** 0.010
RGDPsim	+	0.050*** 0.009	0.035*** 0.011	0.036*** 0.011	0.035*** 0.011
DKL	+	0.203*** 0.043	0.065 0.051	0.059 0.051	0.064 0.051
SQDKL	-	-0.111*** 0.011	-0.062*** 0.012	-0.061*** 0.012	-0.061*** 0.012
DCONT	+	0.516*** 0.039	0.504*** 0.050	0.545*** 0.051	0.389*** 0.047
REMOTE	+	0.518*** 0.095	0.297*** 0.108	0.336*** 0.112	0.293*** 0.108
DROWKL	-	0.001 0.023	0.062** 0.030	0.029 0.030	0.110*** 0.028
Constant		-6.022*** 0.793	-0.640 0.919	-1.556* 0.943	-0.074 0.893
Observations (country-pairs)		15753	15753	15753	15753
Pseudo- R^2		0.229			
σ^2		0.095	1.213	1.221	1.206
Log-likelihood		-4999.999			
Log-likelihood ^a		-76193.787	-62567.977	-62746.380	-62016.239
Used draws from Markov Chain			5000	5000	5000
Thinning ratio ^b			1	1	1
Required number of burn-ins ^b			4	4	3
I -statistic ^b			1.507	1.429	1.233

Notes: Figures below coefficients are standard errors. ^a LeSage (1999a). - ^b Raftery and Lewis (1992a,b, 1995). - ^c The parameter of the χ^2 -distribution of the residuals is set at 4 to account for heteroskedasticity. *, **, *** denotes significance at 10%, 5% and 1%, respectively.

Table 2: Predicted probabilities of a preferential trade agreement (PTA)

Predicted probabilities of a PTA				
Models	Mean	Std.	Minimum	Maximum
Non-spatial probit	0.144	0.166	0.000	0.966
Spatial probit models				
\mathbf{W}	0.107	0.169	0.000	0.996
$\mathbf{W}_{\text{direct}}$	0.093	0.201	0.000	1.000
$\mathbf{W}_{\text{indirect}}$	0.111	0.165	0.000	0.994
Difference in Probabilities to non-spatial probit model				
Models	Mean	Std.	Minimum	Maximum
\mathbf{W}	-0.037	0.070	-0.371	0.407
$\mathbf{W}_{\text{direct}}$	-0.051	0.173	-0.764	0.962
$\mathbf{W}_{\text{indirect}}$	-0.033	0.051	-0.257	0.277

Table 3: Extreme differences of spatial to non-spatial model response probabilities based on the full W-matrix

Country-pair		Non-spatial probit	Difference spatial-non-spatial
Largest negative differences			
Djibouti	Somalia	0.541	-0.371
Oman	Saudi Arabia	0.653	-0.350
India	Iran	0.565	-0.340
Iran	Saudi Arabia	0.644	-0.328
Israel	Saudi Arabia	0.646	-0.327
Largest positive differences			
Aruba	Haiti	0.198	0.407
Bahamas	Haiti	0.260	0.374
Haiti	Netherlands Antilles	0.303	0.352
Belize	Nicaragua	0.535	0.338
Haiti	Nicaragua	0.372	0.306

Table 4: Marginal effects for bilateral economic size and geography variables

Variable	Non-spatial	\mathbf{W}	$\mathbf{W}_{\text{indirect}}$	$\mathbf{W}_{\text{direct}}$
RGDPsum	0.033***	0.016***	0.018***	0.012***
	0.002	0.002	0.002	0.003
NATURAL	0.090***	0.095***	0.095***	0.086***
	0.004	0.008	0.006	0.022
DCONT	0.090***	0.063***	0.071***	0.043***
	0.007	0.008	0.007	0.012
REMOTE	0.090***	0.037***	0.044***	0.032***
	0.017	0.014	0.015	0.013

Notes: Figures below coefficients are standard errors. *** denotes significance at 1%.

Table 5: Sensitivity analysis for the interdependence parameter

Statistic	\mathbf{W}	$\mathbf{W}_{\text{indirect}}$	$\mathbf{W}_{\text{direct}}$
	ρ	ρ_{indirect}	ρ_{direct}
Alternative decay of W-matrix: $w_{ij} = e^{-\text{Distance}_{\ell m}/1000}$			
Mean	0.903***	0.771***	0.961***
Std.	0.025	0.033	0.015
Log-likelihood	-62527.533	-62719.329	-62288.986
Alternative decay of W-matrix: $w_{ij} = e^{-\text{Distance}_{\ell m}/250}$			
Mean	0.844***	0.690***	0.950***
Std.	0.029	0.032	0.018
Log-likelihood	-62354.786	-62612.579	-61831.556
Alternative cut-off level in W-matrix: $e^{-\text{Distance}_{\ell m}/500} = e^{-1000/500}$			
Mean	0.903***	0.770***	0.961***
Std.	0.025	0.033	0.015
Log-likelihood	-62527.533	-62719.897	-62288.986
Alternative level of skewness of χ^2 -distribution of ν_i : homoskedasticity			
Mean	0.748***	0.571***	0.972***
Std.	0.031	0.034	0.015
Log-likelihood	-75948.802	-76134.801	-76083.949
Alternative level of skewness of χ^2 -distribution of ν_i : $r = 2$			
Mean	0.805***	0.635***	0.975***
Std.	0.035	0.039	0.015
Log-likelihood	-62560.941	-62742.691	-62015.795
Alternative level of skewness of χ^2 -distribution of ν_i : $r = 20$			
Mean	0.812***	0.632***	0.978***
Std.	0.035	0.039	0.015
Log-likelihood	-62559.860	-62741.440	-62011.954

Continued on next page

Table 5: Continued from previous page

Statistic	\mathbf{W}	$\mathbf{W}_{\text{indirect}}$	$\mathbf{W}_{\text{direct}}$
	ρ	ρ_{indirect}	ρ_{direct}
Augmenting the specification by political variable			
Mean	0.888***	0.701***	0.986***
Std.	0.037	0.046	0.010
Log-likelihood	-42865.552	-42989.875	-42518.835
Only Custom Unions and Free Trade Areas treated as PTAs			
Mean	0.763***	0.604***	0.959***
Std.	0.035	0.038	0.019
Log-likelihood	-62421.134	-62592.067	-61867.353
Using the same country-pair sample as Baier and Bergstrand (2004)			
Mean	0.774***	0.552***	0.847***
Std.	0.083	0.116	0.059
Log-likelihood	-4321.842	-4340.089	-4307.850

Notes: *** denotes significance at 1.

Table 6: Descriptive statistics

Variable	Mean	Std. Dev.	Minimum	Maximum
PTA	0.144	0.351	0.000	1.000
NATURAL	-8.269	0.778	-9.420	-3.247
RGDPsum	24.681	1.906	18.783	30.182
RGDPsim	-0.733	1.798	-11.000	0.693
DKL	1.847	1.317	0.000	6.100
SQDKL	5.147	6.230	0.000	37.211
DCONT	0.207	0.405	0.000	1.000
REMOTE	8.463	0.165	8.200	9.065
DROWKL	1.605	0.668	0.061	3.884