Spillovers in Trade Agreement Memberships: Does Institutional Homogeneity Matter?†

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Abstract

This paper presents novel evidence that preferential trade agreement (PTA) memberships create spillovers across countries sharing homogeneity in institutions. Using a spatial econometric approach, we take economic system and political regime measured by economic freedom and democracy indexes, respectively, as the fundamental components of institutional distance, and find strong evidence of institutional interdependence both for PTA memberships and for the type of PTA chosen - deep or shallow agreement - based on a sample of 142 countries. These findings are robust to the controlling for the spillovers through geographical proximity as well as possible endogeneity.

Keywords: Preferential trade agreements, institutions, spillovers, spatial autoregressive probit model

JEL Codes: C21, C25, F13, F15, O24

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1 Introduction

In the past decades, the world has witnessed a dramatic growth of the formation of preferential trade agreements (PTAs). However, there is a long-standing debate in academia over its role in multilateralism. On the one hand, some studies support the role of PTAs as stumbling blocks toward global free trade as they may reduce incentives for those involved countries to pursue further multilateral liberalization.[1] On the other hand, other studies reveal that PTAs could be stepping stones toward global free trade, with the "domino theory" of Baldwin (1993) that considers the spillovers in trade agreement memberships as one of the most influential work in this respect.[2]

Specifically, Baldwin (1993) suggests that existing PTAs can change the political balance of the countries that are initially outsiders and thus make them join the existing or establish new PTAs.[3] The theory was not formally tested until Egger and Larch (2008) who are the first to apply a spatial autoregressive probit (SARP) model to confirm the spillovers in PTA memberships through *geographical proximity*.[4] The SARP model is an econometric method that incorporates simultaneity effects among economic decisions, which makes the formal test for spillover effects in PTA memberships feasible and has also been used to explore, for instance, memberships in environmental treaties (see, e.g., Murdoch et al. (2003)).[5]

However, relatively little research has concentrated on the role of trade partners' insti-

[1]See, for instance, Krugman (1989), Bhagwati (1991), Frankel et al. (1995), Krishna (1998), Schiff and Winters (2003), Ornelas (2005), Limão (2007), and Bhagwati (2008).

[2]Other studies suggesting that PTAs are stepping stones toward global free trade include Summers (1991), Estevadeordal et al. (2008), Tabakis and Zanardi (2019), and Kuenzel and Sharma (2021).

[3]See, for instance, Yi (1996), Goyal and Joshi (2006), and Saggi and Yildiz (2010) for relevant theories. Also see Maggi (2014) for an excellent survey on the recent literature on international trade agreements.

[4]See Sapir (2001) as an early empirical study on the domino theory by focusing on the experience of Europe. From a different perspective, Baldwin and Jaimovich (2012) derive an index that measures bilateral trade reliance, and find the self-reinforcing extension of PTAs also depends on such contagion of trade. Chen and Joshi (2010) study the interdependence of free trade agreements (FTAs) by allowing the effect of existing FTAs to vary with the structure of existing FTA relationships.

[5]See, for instance, Case (1992), Case et al. (1993), Beck et al. (2006), Blonigen et al. (2007), Baltagi et al. (2008), Bala et al. (2014), Debarsy et al. (2018), and Zhang et al. (2019) for empirical use of spatial econometrics in multiple research fields.

tutions, particularly *institutional homogeneity*, in the spillovers in PTA memberships as a potential channel, despite that institutions play a significant role in trade agreement formation based on both theoretical and empirical knowledge.[6] Indeed, institutional quality is one of the most important determinants of trade flows toward the formation of PTAs as a source of comparative advantage (see, e.g., Levchenko (2007)).[7] Not only economic institutions such as contract enforcement and property rights as addressed in Levchenko (2007) matter for trade, but also political institutions that of regime type - democracy or not - matter (see, e.g., Mansfield et al. (2000) and Mansfield et al. (2002)), for instance, through improving product quality (see, e.g., Yu (2010)).[8] Not coincidently, both the aspects of institutions are also prominent factors shaping economic growth in the long term (see, e.g., Acemoglu et al. (2005a) and Acemoglu and Johnson (2005)). In spite of the significant contribution made in the literature on the direct impact of institutions on trade, it remains an open question whether institutions have an externality effect through forming a channel of the spillovers in trade agreement memberships as proposed in Baldwin (1993).

Therefore, this research fills into the literature gap by empirically investigating the role of institutional homogeneity - both economic and political institutions - in the spatial spillovers in trade agreement memberships by testing the following two hypotheses: (i) whether PTA memberships are interdependent among countries with homogeneous classification of institutions; (ii) whether such interdependence effect exists in the type of PTA chosen - deep or shallow agreement. To the best of our knowledge, this is the first

[6]The enlargement of the European Union Customs Union (EU) can be a good example as the Copenhagen criteria requires the EU candidate countries to promote their institutions regarding democracy, market economy, among many other aspects to be quantified for formal accession to the EU (see, e.g., Özgüzer and Pensieroso (2013)). Another example comes from the accession of Mexico into North American Free Trade Agreement (NAFTA), as NAFTA is a commitment by Mexico to a future liberalization of economy in exchange for easier access to the markets of the U.S. and Canada (see, e.g., Tornell et al. (1997)).

[7]See, for instance, Dollar and Kraay (2003), Nunn (2007), Costinot (2009), and Chor (2010) for further discussions on institutions as a source of comparative advantage.

[8]See, e.g., Ferrantino (1993), Maskus and Penubarti (1995), Ma et al. (2010), Feenstra et al. (2013), and Sheng and Yang (2016) for more discussions on the relevance of economic institutions in trade, and Morrow et al. (1998), Wong and Yu (2015), and Pinar and Stengos (2021) for the impact of political institutions on trade.

study to test the spillover effects of PTAs through institutional channel and in this sense our paper enhances Baldwin (1993)'s view on domino-like evolution of PTAs and extends Egger and Larch (2008) by considering institutional factor as an additional channel for the interdependence effects of PTA formation.

Our hypotheses are driven by the recent observation that PTAs have become significantly deeper in the new millennium by containing more policy areas and provisions beyond the WTO (see, e.g., Horn et al. (2010), Hofmann et al. (2017), Mattoo et al. (2020)). Compared with shallow provisions that mainly deal with trade-related issues at the border, deep provisions that are mostly behind the border such as competition policy, environment, labour market, intellectual property rights, and state-owned enterprises usually require greater cooperation between trade partners on domestic policies, laws, and institutions, including, for instance, harmonization of regulation and standards, free movement of capital, goods and other factors, restriction of government discretion, amongst others, and as a consequence impose greater limitation on national sovereignty (see, e.g., Feng and Genna (2003), Mansfield et al. (2008), Chauffour and Maur (2011), Vicard (2012), Postnikov (2014), Lechner (2016), and Mattoo et al. (2020)).[9]

As a result, it is as well important for countries to make decision on what specific areas to focus on in their trade policy agenda based on their economic and political context. For instance, democratic countries are found to include more environmental clauses in PTAs than autocratic countries, and on average are more willing to sign deeper PTAs (see, e.g., Mansfield et al. (2008) and Morin et al. (2018)). Meanwhile, upward harmonization to stricter regulation and standards can raise the costs faced by developing countries and may also harm their national objectives, and thus can be resisted by those countries with weak institutions (see, e.g., Chauffour and Maur (2011)). Subsequently, homogeneous countries in institutions tend to have overlapping policy areas in trade agreements that comply with their domestic regulation and standards. Thus, pre-existing PTAs signed

[9]See Maggi and Ossa (2021) for an excellent review on the political economy of deep integration.

by one group of countries, sorted by institutional classification, can work as templates for potential new ones among the same group, lowering both ex-ante negotiation and ex-post implementation costs, for what is known as the *template effect* (see, e.g., Allee and Elsig (2019) and Mattoo et al. (2020)). However, such an externality effect of the PTAs does not apply to other groups of countries that differ in their main settings of institutions. The template effect makes institutional homogeneity a potentially important channel for the spillovers in PTA memberships to test, especially for the decisions on the type of PTAs as the scope of policy in deep PTAs is more reliant on trade partners' institutions.[10]

Fitting a spatial autoregressive probit model (see, e.g., LeSage and Pace (2009) and Martinetti and Geniaux (2017)), our data covers 10,011 country-pairs generated from 142 countries in the year of 2017, with the dependent variable of PTA status obtained from the WTO. To construct the spatial weight matrix that measures the neighbour effect of PTA memberships, we use economic freedom and democracy indexes to calculate institutional distance in a Euclidean form following the literature (see, e.g., Liou and Rao-Nicholson (2017) and Kostova et al. (2020)), since the former one captures the soundness of market economy in multiple aspects including legal system and property rights as emphasized in Levchenko (2007) and the later one captures political regime, both commonly used.[11]

Our results provide strong evidence for the institutional interdependence of PTA memberships as the estimate of the institutional spillover effect is both positive and statistically significant at the 1% significance level. To examine whether the interactive effect driven from institutional similarity is independent of that from geographical proximity suggested by Egger and Larch (2008), we employ the established method of Case et al. (1993) to account for the mixed spillover effect generated from both channels by employing a compound spatial weight matrix, and confirm that institutional channel accounts for around 60% in the optimal compound spatial weight matrix, suggesting the promi-

[10] We construct our hypotheses with more details in Section 2.

[11] Democracy index has been used by, for instance, Mansfield et al. (2000), Mansfield et al. (2002), and Yu (2010). Economic freedom as a measure of institutional quality has been suggested in, for instance, Stroup (2007), Méon and Sekkat (2008), and Francois and Manchin (2013).

nent role of institutions in PTA evolution in an interconnect world.[12] In addition, the spillover effects in PTAs are also found to be quantitatively significant according to the comparison between the direct and indirect marginal effects and an experiment where we evaluate the spillover effects of a one-unit decrease in the estimated minimum utility gains from forming PTAs for selected countries and their trade partners.

We next examine whether the spillovers exist in the decisions on the type of PTAs - deep or shallow agreement - through institutional homogeneity. As in Mattoo et al. (2020), we define those PTAs involving more than 15% of all the available provisions as deep PTAs and the rest as shallow ones.[13] The estimation results confirm the hypothesis that institutional homogeneity plays a key role in the spillovers in the type of PTA memberships, which is even more prominent than in the spillovers in PTA memberships, as the estimated coefficient of the spatial lag term is double in magnitude from the previous one and it is still statistically significant. Allowing for the mixed spillover effects through both institutional homogeneity and geographical proximity, we find the optimal spatial weights generated from institutional channel is around 90%, suggesting that the institutional channel dominates the spillovers in the type of PTAs. The spillover effects are also found to be quantitatively important.

Our empirical findings hold up well in various robustness checks. First, it is necessary to consider possible endogeneity as the causality between PTA formation and trade partners' institutions can run two-ways.[14] We thus employ a new estimation procedure allowing for endogeneity of both the spatial weights and the explanatory variables following Qu and Lee (2015), and confirm that our main results continue to hold with the

[12]The optimal compound spatial weight matrix is decided by the model through comparing the goodness of fit measured by negative log-likelihood function.

[13]Our findings are not qualitatively changed by using other threshold values to define deep PTAs such as 25%.

[14]On the one hand, institutions can affect PTA formation through either direct or indirect channels such as neighbour effects. On the other hand, trade partners' institutions can be improved during the negotiation of PTAs so as to meet the accession criteria. See, for instance, Acemoglu et al. (2005b), Levchenko (2013), and Stefanadis (2010) for the impact of trade on institutions.

correction for potential endogeneity.[15] Second, we take all the EU countries as a single entity in the empirical analysis to examine the generality of our results, as a significant part of preferential relationships is built between the EU countries, and we confirm that our findings still hold true.

To sum up, exploring the spillover effects of PTAs based on a novel spatial econometric approach, this paper contributes to the literature of trade agreement determinants (see, e.g., Mansfield et al. (2002), Baier and Bergstrand (2004), Endoh (2006), Liu (2008), Mansfield et al. (2008), Vicard (2012), Liu and Ornelas (2014), and Davis and Wilf (2017)). Particularly, showing that *institutional homogeneity* makes up an additional channel for the spillovers in both PTA memberships and their type, our paper complements previous empirical studies on the spillover effects of PTA memberships suggested by Baldwin (1993) through different channels, for instance, Egger and Larch (2008) through *geographical proximity* and Baldwin and Jaimovich (2012) through *trade reliance*.

The reminder of the paper is organised as follows. In Section 2, we proceed to formally define the institutional spillover hypotheses that we would test in this paper. Section 3 presents data and empirical specification. Section 4 shows the estimation results and some robustness checks. In Section 5, we conclude.

2 Institutional Spillover Hypotheses

According to the principles and the relevant chapters of the WTO agreement, specifically GATT Article XXIV, WTO membership countries are allowed to form customs unions or free trade agreements where two or more countries will decide to completely eliminate all tariffs between each other, while, without eliminating tariffs on goods imported from the rest of the world.

Since the world turned to the 21st century, we have observed dramatic emergence of

[15] We also conduct a small set of Monte Carlo simulations to assess the finite sample performance of the proposed estimator.

deep PTAs, which refer to those trade agreements with a large set of deep provisions beyond the traditional scope and usually behind the border that include but not limited to the following significant ones: competition policy, investment, services, subsidies, immigration, environment, labor, intellectual property rights, state-owed enterprises. Using the deep trade agreement database constructed by Hofmann et al. (2017) and Mattoo et al. (2020), Figure 1 takes the birds-eye view of the evolution of the PTA depth measured by the total number of policy areas and provisions starting from 1960s, and shows a clear rising trend of the PTA depth over the last two decades.

As deep PTAs involve trade partners' domestic laws, trade partners' institutional background such as economic system and political regime as addressed in the trade literature become a potentially important determinant of not only PTA formation but also its content in terms of depth. For instance, the EU countries tend to sign deeper PTAs, compared with countries in other regions, as most of them have a stable democracy and a functioning market economy (see, e.g., Horn et al. (2010)).

To offer a quantifying picture between the depth of PTAs and trade partners' institutional background, we measure the PTA depth by either the number of policy areas or provisions, and we follow the literature mentioned in the previous section to take democracy and economic freedom as the representative aspects of institutions, with the data of democracy index obtained from Polity IV Project and that of economic freedom index obtained from the Fraser Institute, where larger values indicate higher scores toward consolidated democracy and complete economic freedom, respectively. Letting $PIVsum$ and $EFWsum$ be the aggregate country-pair scores for democracy and economic freedom, respectively, we report the Pearson correlation coefficients between the two institutional scores with the PTA depth in Table 1. As shown in Table 1, the Pearson correlation coefficients are all significantly positive, which implies that trade partners with better institutional quality in terms of democracy and market economy, ceteris paribus, tend to sign deeper PTAs.

To further ensure that the observed relationships between the PTA depth and institutional scores are not driven by other characteristics of countries, we carry out a number of OLS regressions and take the results as supplemental evidence, with the two measures of the PTA depth (i.e., the number of areas and provisions) as the dependent variable, separately, and the two institutional variables $PIVsum$ and $EFWsum$ as key independent variables, along with a set of other control variables as potential determinants of PTA status following the literature (see, e.g., Baier and Bergstrand (2004)).[16] The results are shown in Table 2, and we can see that the estimates of the effects of $PIVsum$ and $EFWsum$ on the PTA depth remain positive and statistically significant regardless which measure of the PTA depth is used. These findings are consistent with Vicard (2012), where the author finds that institutional determinants including the degree of democracy are more relevant to deep PTAs than to shallow ones and further suggests that similar domestic institutions such as political regime and governmental type can make it more possible for a pair of countries to give up some national sovereignty and transfer it to supranational organizations such as deep PTAs.[17] Overall, trade partners with a higher degree of democracy and a better market economy tend to concentrate on a larger scope of policy with many behind-the-border issues in their trade agreements.

On the other hand, the non-trade articles raised by countries with high degree of democracy and economic freedom are not equally acceptable by countries with low institutional scores. It is partially because those countries lack both domestic laws and enforcement powers to execute the horizontal commitments made in PTAs, and furthermore it might even not be in their best interest to adopt such commitments in the first place as, for instance, low standards of environmental protection may distort competition and thus strengthen export competitiveness (see, e.g., Esty and Geradin (1997) and Chauffour and Maur (2011)).

[16] We will introduce the control variables with more details in Section 3.1.

[17] Note that the definition of deep PTAs in Vicard (2012) is in a more aggregate level compared with ours which stands for customs unions and common markets. Also, Vicard (2012) focuses on institutional effect on PTA formation, while we focus on the spillover effects of PTA formation.

Since country-pairs with more homogeneous institutions of economy and politics tend to share closer views on what to negotiate and include in their trade agreements, pre-existing PTAs can thus be more useful templates for potential PTAs among country-pairs sharing homogeneous institutions by reducing both the ex-ante negotiation and ex-post regulatory costs for potential new ones, the so called *template effect*. These observations motivate us to further explore possible institutional interdependence of PTA formation.

For instance, due to the pre-existing trade agreement reached by Canada and the U.S. (i.e., NAFTA), a relatively deep one that covers a large set of topics other than trade facilitated ones such as environment, labor, intellectual property rights, state-owned enterprises, corruption, as well as the pre-existence of the EU Customs Union which deemed to be the so far deepest economic integration, the negotiators of the potential trade agreement between Canada and the EU (i.e., CETA) could hold more consensus on what to discuss and include in their new trade agreement by taking both NAFTA and the EU Customs Union as templates so as to facilitate the negotiation process.[18] In addition to the template effect on negotiations, from the Canadian point of view, the overlaps between NAFTA and CETA make it less costly to revise domestic laws for the implementation of CETA as well as to set up a new regulatory agency for the monitoring and enforcement of CETA since the Canadian authorities have already had previous experience in that respect. However, the pre-existing deep PTAs achieved by countries with established institutions can only have very limited impact as templates on potential trade agreements between countries with weak institutions, because, as documented, their policy targets of PTA formation are different, conditional on their domestic institutions.

The template effect in the formation of PTAs is not entirely new to the literature, for instance, it has also been suggested by Mattoo et al. (2020). Although their focus is not on the interdependence effect of PTA formation, they construct an index of similarity for

[18]Issues related to environment and labor market are put separately in the North America Agreement on Labor Cooperation and North American Agreement on Environmental Cooperation as supplemental materials to the NAFTA.

PTAs based on the ratio between the number of policy provisions included in two agreements, and simply document that PTAs are highly similar within each trading bloc due to the template effect of the pre-existing PTAs. Based on text analysis, Allee and Elsig (2019) also find that PTA contents are heavily copied and pasted from the pre-existing templates, which holds true for both PTAs concluded by developing and developed countries.

We conjecture that the template effect may drive the spillovers of PTA formation among countries with similar institutional background. Such logic also applies to the choices of PTA type that of deep or shallow agreement, as the template effect generated from the pre-existing deep PTAs could encourage trade partners with similar institutions to choose deep PTAs over shallow ones. In fact, the interactive effects in the decisions on the type of PTAs can even be more prominent than in the formation of PTA memberships, since deep PTAs are more relevant to domestic institutions as remarked earlier. We therefore aim to test the following PTA formation spillover hypotheses:

Hypothesis I: PTA memberships are interdependent based on institutional homogeneity between country-pairs, and such an interdependence effect decreases with the institutional differential.

Hypothesis II: The decisions on the type of PTAs - deep or shallow agreement - are also interdependent through the institutional channel, which can be more prominent than the interdependence of PTA memberships.

3 Data and Empirical Strategy

In this section, we set up the empirical strategy to test the institutional spillover hypotheses of both PTA memberships and their type defined in Section 2. We begin with presenting the variables and data we use in Section 3.1 and in Section 3.2 we introduce the benchmark specification based on a spatial autoregressive probit model and explain the construction of spatial weights that capture the interdependence of PTAs through institutional similarity.

3.1 Data

In this study, we follow the most common definition of PTAs used in the academic literature with the original data of PTA status collected from the WTO by Egger and Larch (2008), including all the bilateral and regional trade agreements notified to the WTO from 1950 to 2017. As the dataset of PTA status does not include the information of depth such as policy areas and provisions, we thus supplement the dataset with deep trade agreement data obtained from the two recent datasets constructed by Hofmann et al. (2017) and Mattoo et al. (2020). Although both datasets address PTA depth, their focuses are slightly different. Particularly, Hofmann et al. (2017), as an earlier attempt, investigate the coverage of policy in PTAs at an area level, while Mattoo et al. (2020) build up their dataset based on provision-level policy coverage. There are, in total, 52 policy areas and 937 policy provisions documented in the two datasets, with the actual number of areas or provisions measuring PTA depth.

As noted earlier, we consider political regime and economic system as the key characteristics of domestic institutions and take democracy index and economic freedom index to measure them as both institutional aspects are found to be the key determinants of growth and trade toward the formation of PTAs, from which we calculate the institutional distance and construct the institutional spatial weight matrix that measures interactive effects between country-pairs. The index of democracy is obtained from the Polity IV Project, and the index of economic freedom is obtained from the Fraser Institute, and both have been popularly used in the academic literature to measure countries' institutions (see, e.g., Mansfield et al. (2000), Mansfield et al. (2002), Stroup (2007), Yu (2010), and Francois and Manchin (2013)). For both the indexes, larger values represent higher scores toward consolidated democracy and complete freedom of economy, respectively. Specifically, the polity score measures the regime authority spectrum on a scale ranging from

-10 for hereditary monarchy to 10 for consolidated democracy. As a relatively aggregate index evaluating the soundness of market economy that could be relevant to trade, economic freedom is calculated from five areas, including size of government, legal system and property rights, sound money such as inflation management, freedom to trade internationally, and regulation (or restriction), with its highest score at 8.90 for Hong Kong and lowest score at 2.72 for Venezuela in the year of 2017. The construction of institutional distance and corresponding spatial weights based on the original data is presented in Section 3.2.

In line with the established literature on trade agreement formation such as Baier and Bergstrand (2004) and Egger and Larch (2008), we include the following control variables in our baseline regressions: the log of the inverse of the great circle distance between the capitals of trade partners, denoted as $NATURAL$, the remoteness of a pair of continental trade partners from the rest of the world, denoted as $REMOTE$, the total bilateral market size measured by total GDP of trade partners, denoted as $GDPsum$, the similarity of trade partners in terms of GDP scale, denoted as $GDPsim$, the absolute difference in GDP per capita between trade partners, denoted as DKL and its square $SQDKL$, and the relative factor endowment difference between the rest of the world and a given pair of trade partners, denoted as $DROWKL$.

The geographical data including latitude and longitude coordinates of national capitals is taken from the GeoDist dataset of CEPII to calculate the geographical distance between countries such as $NATURAL$ and $REMOTE$. In addition, we also use the GeoDist dataset to calculate the average distance between country-pairs to construct geographical spatial weights, as geographical vicinity can be another source of spillover effects for PTA memberships. Other variables are retrieved from the WDI database of the World Bank. Finally, we also control for the absolute differential of institutional scores between trade partners, $PIVdist$ and $EFWdist$, as their effects might be correlated with the institutional spatial weight matrix and excluding the two variables can potentially make the estimate of the latter upward biased. Specifically, trade partners with heterogeneous institutions may be less likely to form PTAs, resulting in countries with high scores of institutional indicators forming PTAs with each other and countries with low scores also forming PTAs with each other (as their institutional heterogeneity is small), something that may be misattributed to the effect drawn from institutional spillovers in PTA memberships and their type.

The final dataset we use in the main regressions is taken from the year of 2017 and covers a sample of 142 countries and 10,011 country-pairs. The summary statistics of the key variables is presented in Table 3, and the countries in our sample are listed in the appendix. For the benefit of robustness check, we also construct a balanced panel dataset for the period from 1996 to 2017 in four five-year intervals, with 111 countries included in total, and the summary statistics of the panel data is also shown in the appendix.

3.2 Empirical strategy

3.2.1 Specification

We now proceed to test the spillover effects of PTA memberships and their type through institutional homogeneity between country-pairs. We find that spatial autoregressive probit models (or spatial autoregressive binary choice models) are well-fitted to capture interactive effects of trade partners. Spatial econometric methods, especially those with limited dependent variable (e.g., binary or multinomial) as what we will follow in this study, have been found increasingly useful in various fields of economic research because of its advantages in addressing simultaneity effects in economic activities.

For instance, Case (1992), as an early attempt, uses a SARP model to examine the spillover effect in the adoption of new technology in Indonesia. Applying to housing economics, Bala et al. (2014) and Zhang et al. (2019), along with a number of other work, use spatial econometric models to investigate the interdependence of real estate markets across regions. In the topics of international economics, Blonigen et al. (2007) use a spatial autoregressive model to check the spatial interdependence in FDI activities, while Baltagi et al. (2008) study the effect of regional trade agreement on FDI by using a spatial heteroscedasticity and autocorrelation consistent estimator. Allowing for spillovers across financial markets, a recent paper by Debarsy et al. (2018) addresses the spatial dependence of sovereign bond spreads. In addition, spatial econometrics is also found useful in correcting potential upward biases in the estimation of border effect (see, e.g., Behrens et al. (2012)). More relevant to this paper, focusing on the memberships of international organizations, Murdoch et al. (2003) and Egger and Larch (2008) employ the methodology to study the simultaneity in the formation of environmental treaties as well as preferential trade agreements.

The above empirical papers show us how spatial econometrics can improve traditional estimation by allowing for interdependence among economic decisions when agents can learn from each other as in this study through the template effect for decisions on trade agreement formation. In line with the literature, particularly LeSage and Pace (2009), we set up our spatial autoregressive probit model in matrix form as follows to test the first hypothesis about PTA memberships:

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

where PTA^* and PTA are $n \times 1$ vectors, 0 is an $n \times 1$ vector of zeros, W^I is an $n \times n$ spatial weight matrix constructed from country-pair institutional differential, X is an $n \times k$ matrix of other explanatory variables, including $NATURAL$, $REMOTE$, $GDPsum$, $GDPsim$, DKL, $SQDKL$, $DROWKL$, $PIVdist$, $EFWdist$, and also a constant term. $I[A]$ is an indicator function, taking a value of one if event A occurs and zero otherwise. ρ is the so-called spatial lag parameter and $\rho = 0$ implies the absence of interdependence of PTA memberships, while $\rho > 0$ implies the existence of interdependence of PTA memberships with its effect decreasing with the institutional differential between country-pairs. Thus, a significantly positive estimate of ρ would verify the first institutional spillover hypothesis we put forward in Section 2. β is a vector of unknown parameters of other explanatory variables, ε is an $n \times 1$ vector of errors, and finally, n equals the number of country-pairs (or observations). Specifically, each element of the vector PTA^* can be read as the minimum value of utility differential for a pair of trade partners between having or not having a bilateral PTA. For instance, for country-pair i with trade partners i_1 and i_2, $PTA_i^* = min(\Delta U_{i_1}, \Delta U_{i_2})$, where ΔU_{i_1} and ΔU_{i_2} denote the utility differences for country i_1 and country i_2 between signing or not signing a bilateral PTA (or joining or not joining a regional PTA), respectively. In the data, PTA_i^* is not directly observable, and instead, we can only observe PTA_i as the output of the indicator function $I[PTA_i^* > 0]$, which equals 1 if and only if both countries i_1 and i_2 are better off in terms of welfare by having a preferential trade relation between each other. Thus, PTA_i is a binary variable which takes value 1 if there is a preferential trade relation within a country-pair and takes value 0 otherwise.

In testing the second hypothesis about the type of PTAs, we define the dependent variables PTA^* and PTA in Equation 1 in the way that the new PTA equals 1 if there is a deep PTA and 0 if there is a shallow PTA, and the new PTA^* can be read as the minimum value of utility differential for a pair of trade partners between having a deep or shallow PTA. We note that the sample size is also smaller in this case as only trade partners with agreements formed are considered.

3.2.2 Spatial weights

Next, we proceed to set up the institutional spatial weight matrix W^I determined by institutional differential, whose elements measure the interdependence effect of PTAs through institutional homogeneity between country-pairs. As suggested by Beck et al. (2006), space is more than geography in the context of spatial econometrics, especially when it comes to political economy, and institutional distance that we consider in this study is perhaps one of the most important dimensions other than geography that can affect international business (see, e.g., Liou and Rao-Nicholson (2017) and Kostova et al. (2020)). Thus, we follow the literature and use institutional differential to construct spatial weights so as to test the institutional spillovers in PTA memberships as well as their type. Specifically, a typical spatial weight matrix is row-normalized and has all its diagonal elements being zeros. Suppose that one country-pair i consists of two countries i_1 and i_2, and another country-pair j consists of two countries j_1 and j_2. Following the standard way in the literature especially in international business studies, we use Euclidean distance calculated with the standardized values of two institutional variables to measure the institutional differential between country-pairs i and j (see, e.g., Liou and Rao-Nicholson (2017) and Kostova et al. (2020)):

$$Distance^{I}_{ij} = [(d_i - d_j)^2 + (e_i - e_j)^2]^{\frac{1}{2}}, \tag{2}$$

where d_i and d_j represent the aggregate scores of democracy for country-pairs i and j, respectively; e_i and e_j represent the aggregate scores of economic freedom for country-pairs i and j, respectively. Specifically, $d_i = d_{i_1} + d_{i_2}$ and $e_i = e_{i_1} + e_{i_2}$, where d_{i_1} and d_{i_2} stand for the individual levels of democracy for country i_1 and country i_2, respectively, and e_{i_1} and e_{i_2} stand for the individual levels of economic freedom for country i_1 and country i_2, respectively. Since the two aggregate scores of institutions have very different ranges in the raw data, they are both standardized so as to measure how far from the mean a data point is, before they are used to calculate the Euclidean distance. Next, we construct the spatial weights from a negative exponential function, one of the most popularly used weight functions in the spatial econometric literature:

$$w^{I*}_{ij} = e^{-Distance^{I}_{ij}}, i \neq j, \tag{3}$$

which denotes the $(i,j)^{th}$ element in the negative exponential distance weight matrix, with all the diagonal entries set to zero.[19] Due to the memory limitations of a standard personal computer, we replace the cells of the spatial weight matrix that are less than the value of the 95th percentile with zeros. As a result, there are 5% of the cells in the spatial weight matrix are non-zero.[20] As the final step to obtain the spatial weight matrix ready for regressions, we row-standardize the above negative exponential distance weight matrix as follows:

$$w_{ij}^I = {w_{ij}^I}^* / \sum_j {w_{ij}^I}^*, \tag{4}$$

which denotes the $(i,j)^{th}$ element in the final spatial weight matrix W^I, with the sum of each row equaling to one.

After formalizing the SARP model and the corresponding spatial weight matrix, we move to estimation. Due to the shortcomings of Bayesian methods in spatial probit regressions including issues of accuracy and the extreme computational burden for large samples, we adopt the approximated maximum likelihood estimation method introduced in Martinetti and Geniaux (2017), where the vector of the regression errors are assumed to follow a multivariate normal distribution whose parameters depend on the spatial structure of the observations. Thus, our estimation procedure is unlike the previous studies on the interdependence effect of PTAs that rely on Bayesian methods. According to a variety of simulation studies, the adopted approximated maximum likelihood method achieves impressive reduction in computation time and in the meanwhile performs relatively better than other methods in terms of accuracy.[21] Note that the fast calculation is of an utmost important concern for us as we have a fairly large spatial weight matrix of dimension $10,011 \times 10,011$.

[19] Another popular form of spatial weights is inverse distance weighting, and we use it in a robustness check and find that our main findings are insensitive to the choice of the spatial weight functions. Due to space limitation, we did not report the corresponding results in the paper, which is available from the authors upon request.

[20] The choice of the threshold value here for the benefits of computation does not affect our main results.

[21] For instance, Bivand and Piras (2015) present a comprehensive comparison between different implementations of estimation for spatial econometrics.

4 Results

Section 4 contains three subsections. First, we examine the institutional spillovers in PTA memberships in Section 4.1. In Section 4.2, we distinguish deep PTAs from shallow PTAs, so that we investigate whether decisions on the type of PTAs are interdependent through institutional homogeneity. Finally in Section 4.3, we conduct two robustness checks with the first one dealing with possible endogeneity and the second one considering the EU member states as a single entity.

4.1 Spillovers in PTA memberships

4.1.1 Results

We start by discussing the empirical results for the institutional spillovers in PTA memberships. As noted, all specifications include the standard determinants of PTA status used in the literature as control variables. In column (1) of Table 4, we include the theoretical predictions of the sign of each explanatory variable to examine whether the estimates from our specifications are consistent with those predicted by the relevant theories. Next to that, column (2) presents the results from a standard probit model that does not allow for spatial interdependence. Column (3) presents the key results estimated from the SARP model in Equation 1 that allows for the spillover effect of PTA status generated from the institutional similarity between countries. Finally, the specification used in column (4) examines the interdependence effect through geographical proximity, by replacing the institutional spatial weight matrix with the geographical one following the earlier literature.

The baseline results in column (3) of Table 4 present strong evidence verifying the institutional spillover hypothesis of PTA memberships as proposed in Section 2. The spatial lag term of interest, $W^I \cdot PTA^*$, has a positive estimate of coefficient ρ at the 1% significance level. The estimate along with the statistical test indicate that the decisions on forming PTAs interdepend on each other across trade partners, and according to the nature of the institutional spatial weights, such an interdependence effect decreases with the institutional distance between two country-pairs as the template effect of the pre-existing PTAs on potential PTAs is weaker when institutional differential is larger, consistent with what we have documented in Section 2. In other words, if two country-pairs share similar institutions in terms of political regime and economic system, their decisions on whether or not to join or establish trade agreements will be significantly correlated. By contrast, if one pair of trade partners has, for instance, a set of liberal institutions with consolidated democracy and complete economy freedom such as electoral system and unobstructed competition, while the other pair of trade partners has its domestic institutions with relatively more centralized government and also more government intervention in economy, their decisions on PTA formation will be not that inter-related, given that their policy targets to achieve in PTAs are relatively different.

Turning to the control variables, as presented in columns (2)-(4), all the estimates of the standard control variables have the signs in agreement with the theoretical predictions, although their values are slightly different across specifications. Specifically, both total bilateral market size and market size similarity, denoted by $GDPsum$ and $GDPsim$, respectively, always have positive effects on the probability of forming bilateral PTAs. In addition, the difference of development stage measured by the absolute difference of GDP per capita also positively contributes to the formation of PTAs, although its impact is nonlinear as the square of the variable has a significantly negative effect on the formation of PTAs. Consistent with the previous findings, country-pairs far away fom the rest of the world are more likely to form PTAs, indicated by the estimated coefficient of the remoteness of country-pairs, while country-pairs are less likely to form PTAs if their relative factor endowment is more different from that of the rest of the world, indicated by the estimated effect of $DROWKL$.

Regarding the estimated coefficients of another two controlled institutional variables:

the absolute difference of democracy score $PIVdist$ and the absolute difference of economic freedom score $EFWdist$, we find that both coefficients have the expected negative signs and are highly statistically significant across all the specifications, which implies that trade partners experiencing a smaller institutional differential with regard to political regime and economic system tend to have a better chance to form a PTA with each other. Comparing the coefficients in column (2) without spillovers and column (3) with institutional spillovers, we find that there is a big drop for both the coefficients in front of $PIVdist$ and $EFWdist$ when we allow for institutional interdependence effect in the model, suggesting that the omitted institutional interdependence effect could make other estimates biased, especially those institution-related ones as the omitted variable effect can be misattributed to them.[22]

The interdependence effect of PTA memberships could be not only driven by institutional similarity proposed in this study, but also geographical vicinity, as suggested by the previous study. Therefore, we estimate the SARP specification with a geographical spatial weight matrix (see, e.g., Egger and Larch (2008)).[23] As shown in column (4) of Table 4, our results confirm the interdependence of PTA memberships based on geographical vicinity. Similarly, we find that the coefficient estimates of the geography-related variables such as $NATURAL$ and $REMOTE$ are smaller in this specification than those in columns (2) and (3), again indicating the importance of allowing for interdependence effects in modelling trade agreement memberships, which otherwise will make other estimates biased.

Next, it is natural to think about accounting for both the spatial effects driven from institutional homogeneity and geographical proximity in a model so that we can examine whether the institutional spillover effect of PTA memberships is robust to the inclusion of the geographical spillover effect of PTA memberships. Therefore, we revise the SARP

[22]In another robustness check, we run a SARP regression without controlling for $PIVdist$ and $EFWdist$, and find that the estimated value of ρ becomes larger in this case, indicating larger effect of institutional interdependence, consistent with our expectation that $PIVdist$ and $EFWdist$ could be positively correlated with the institutional spatial weights and should be controlled for.

[23]Please see the appendix for details about the construction of geographical spatial weights.

model by allowing for a compound spatial weight matrix that accounts for institutional and geographical distances simultaneously. Specifically, we have $W^C = \alpha W^G + (1 - \alpha)W^I$, where W^G is the spatial weight matrix constructed from countries' geographical distance based on the coordinates of national capitals, and $\alpha \in [0, 1]$ denotes the weight of geographical spatial weight matrix in the compound spatial weight matrix (see, e.g., Case et al. (1993) and Zhang et al. (2019)). As α can be any value ranging from 0 to 1, we compare the goodness of fit measured by log-likelihood across specifications using different values of α to find the optimal one that minimizes the negative log-likelihood function.[24]

As shown in Table 5, the estimation results based on different values of the compound spatial weight α ranging from 0.1 to 0.9 with an equal increment of 0.1, along with the related negative log-likelihood, suggest a U-curve relationship between α and the negative log-likelihood, and further suggest that the minimum value of the negative log-likelihood is achieved when α equals 0.4. This is confirmed by further specifying α from 0.35 to 0.45 with an equal 0.01 increment. The optimal value of α decided by the model suggests that 60% of the spillover effects of PTA evolution are driven by the homogeneity of domestic institutions, while the rest 40% are driven by the geographical vicinity. Reading column (5) for the case of $\alpha = 0.4$, we find that the estimated coefficient of the mixed spillover effect is between the estimates of the pure spillover effects generated from institutional similarity and geographical vicinity as shown in Table 4, with its significance level at 1%, confirming the existence of such mixed spillover effect in PTA formation. The estimated direct effects of institutional heterogeneity within a pair of trade partners measured by $PIVdist$ and $EFWdist$ are still smaller than those without controlling for institutional interdependence. Similar pattens are observed for bilateral distance ($NATURAL$) and remoteness ($REMOTE$) of a pair of trade partners.

For further robustness, we also present the results drawn from a panel data specifica-

[24]Please see the appendix for details about the compound spatial weight matrix.

tion in Table 6.[25] The balanced panel dataset covers 1996 to 2017 in four five-year intervals, with 111 countries, in total, included. Note that a panel data model uses information on the dynamics of variables and can thus make short-to-medium-term implications, while a cross-sectional model emphasizes the long-term implications. Column (1) presents the estimation results without any spatial effects, and columns (2) and (3) present the results with the spatial weight matrices of institutional homogeneity and geographical vicinity, respectively. First, we observe a highly significant positive estimate for the coefficient of the spatial lag term in column (2) at the 1% significance level, which provides additional evidence for the hypothesis of the institutional spillovers in PTA memberships. Second, the results in column (3) also confirm the existence of the neighbour effect through geographical vicinity. Therefore, we include both the neighbour effects generated from institutional similarity and geographical proximity, and we show the results in column (4). After including geographical channel, the institutional spillover effect measured by the estimated coefficient of the institutional spatial lag term becomes a little smaller (from 2.138 to 1.854), while remaining statistically significant, confirming the robustness of our findings. Finally, we find that all the estimates of the control variables have the signs in agreement with the previous theoretical predictions in columns (1)-(3), although the effect of the relative factor endowment difference measured by $DROWKL$ is not significant in columns (3) and (4). Overall, the results based on panel data are consistent with our previous findings and confirm the first hypothesis.

4.1.2 Quantification

As the next step, we make inference from the SARP model with the optimal compound spatial weight matrix based on the results in Table 5. According to the nature of probit models, the coefficient estimates presented do not stand for the differences in probability, but for the shifts in the cumulative distribution function of the standard normal

[25]Please see the appendix for details about the panel data specification

random variable, which makes it not straightforward to quantify either the direct effect of each variable or the spillover effect on the probability by simply using the coefficient estimates. Thus, we calculate the marginal effects of the SARP model with the compound spatial weight matrix to directly quantify the impact of different factors on the expected probability of forming PTAs.

As shown in Table 10, the marginal effects can be distinguished between direct and indirect ones, which is different from standard probit models, with the sum of them called total effects.[26] By definition, direct effects measure the average over all the country-pairs of the marginal effect of an explanatory variable of a single country-pair on the choice probability of the same country-pair. By contrast, indirect effects represent the average over all the country-pairs of the marginal effect of an explanatory variable of a single country-pair on the choice probability of all other country-pairs. The indirect effects are generally at par with the direct effects for most of the variables in the top half of Table 10, illustrating the substantial role played by the spillovers in the evolution of PTAs through institutional and geographical proximity. For instance, a 10 percent increase in the bilateral market size for a pair of trade partners measured by $GDPsum$ can increase the probabiliy of forming PTAs by around 0.8 percent in total, with less than 0.5 percent increase for the same pair of trade partners and more than 0.3 percent increase for its institutional and geographical neighbours.

Motivated by the recent concerns about deglobalization in some countries (see, e.g., James (2018)), we next allow for two shocks of deglobalization to the expected utility gains from forming PTAs captured by the latent dependent variable (PTA^*) for selected countries, so as to better understand the interdependent transmission of PTA formation in terms of its magnitude and also to complement the analysis of the marginal effects explained above. In the first experiment, we generate a one unit negative shock to the errors in the latent model (1) so as to negatively affect PTA^* for G7 countries: Canada,

[26]Please see the appendix for details about the inference of SARP model.

France, Germany, Italy, Japan, the United Kingdom, the United States, and their trade partners, and examine its impact on the probabilities of forming PTAs for both themselves and other country-pairs, where PTA^* can be understood as the minimum value of the expected utility gains for a pair of trade partners to sign PTAs between each other. If the institutional spillover channel studied in this paper is quantitatively important, we expect the shock to have large cross-country impact on the probabilities of forming PTAs.

First, we find the average decrease in probabilities of having PTAs caused by the shock is 4.2% for all country-pairs in the sample. To compare the changes between country-pairs with at least one G7 country, denoted as G7 country-pairs, and those without any G7 country, denoted as non-G7 country-pairs, we show the distributions of probability changes separately for the two groups in Figure 2. Intuitively, since the shock directly affects the expected utility gains from forming PTAs for G7 countries and their trade partners, G7 country-pairs naturally experience larger decreases in their probabilities of forming PTAs, with the average decrease equal to 27.7%. More importantly, due to the spillover effects through both institutional and geographical channels, there is a significant number of non-G7 trade partners experiencing large decreases in their probabilities of forming PTAs as shown in Figure 2, with the largest drop equal to 10.5% for Costa Rica-Sri Lanka and the the average decrease equal to 1.6%.

We repeat the above experiment for another group of countries that consists of China, France, Russian Federation, the United Kingdom, and the United States or the five permanent members of the United Nations Security Council, denoted as P5 countries for simplicity, as they are also regarded as both economically and politically influential countries and cover a wider range of political and economic institutions than G7 countries according to the indexes of democracy and economic freedom. Generally, we find similar patterns that highlight the quantitative significance of spillover effects of PTAs especially through institutional homogeneity. Specifically, We find that the average decrease in probabilities of having trade agreements equals 2.9% in the full sample or, in more disaggregated level, 26.3% and 1.2% for P5 and non-P5 country-pairs, respectively. The distributions of probability changes for P5 and non-P5 country-pairs are shown in Figure 2, separately.

4.2 Spillovers in the type of PTAs

4.2.1 Results

After examining the first hypothesis of PTA memberships, we shift the focus to the type of PTA memberships to further examine whether such spillover effect exists in the decisions on forming deep or shallow agreements. The baseline results are shown in Table 7, where column (1) presents the results without spatial linkages, column (2) presents the results with institutional spatial linkage of our interest, and finally column (3) presents the results with geographical spatial linkage.

Interestingly, the results in column (2) suggest that the spillovers in the type of PTAs through institutional similarity is not only statistically significant, but also economically more important than its counterpart measuring the spillovers of PTA memberships, as the estimated coefficient of the spatial lag term ($W^I \cdot PTA^*$) is 0.625, which is double of the previous one in column (2) of Table 4. This finding is consistent with the fact as discussed in Section 2 that deep PTAs are more dependent on domestic laws and policies than shallow ones as they deal with more behind-the-border issues, and thus decisions on forming deep or shallow PTAs are more relevant to trade partners' institutional background. In this sense, if a country-pair chooses to pursue a deep PTA, its neighbour countries measured by institutional homogeneity will be more likely to follow the path to signing a deep PTA rather than a shallow one. After allowing for institutional interdependence in the model, the estimates for the direct effects of institutional distance between a pair of trade partners ($PIVdist$ and $EFWdist$) become much smaller, suggesting the importance of allowing for spillovers in predicting the type of PTAs as ignoring them can make other estimates biased. In column (3), a similar effect is found by considering geographical channel, indicating that the decisions on the type of PTAs also depend on each other through geographical proximity.

Now we consider the control variables as we want to know what country-pair characteristics are relevant to the decisions on forming a deep or shallow trade agreement. Similar to the results of PTA membership determinants, in Table 7, we find that $NATURAL$, $GDPsum$, and $GDPsim$ all have positive impact on the choice of deep PTAs, with statistical significance at 1% across specifications, indicating that trade partners closer to each other, having larger and more similar markets are more likely to integrate with each other in a deep way by forming comprehensive trade agreements with a large policy focus. Also, the relative factor endowment difference between a pair of trade partners and the rest of the world measured by $DROWKL$ has very similar estimated coefficients to the ones in Table 4, which positively contributes to the choice of deep PTAs. Again, DKL is found to have a nonlinear effect on the preferences over deep PTAs, with the estimate of DKL, for instance, equal to 0.531 as in column (1) drawn from a simple probit model and that of $SQDKL$ equal to -0.102. Moreover, both values are much larger than their counterparts in the results of PTA memberships in absolute terms, suggesting the absolute difference in GDP per capita is more relevant to the decisions on the content of PTAs than the simple status of PTAs.

One unexpected result comes from $REMOTE$, whose estimate is negative in the prediction of deep agreements, but not always significant as in Table 7, which means trade partners further away from the rest of the world do not tend to prefer deep agreements over shallow agreements, although they are found to be more likely to form PTAs with each other in general as shown in Table 4. Finally, consistent with prior expectations, differentials in domestic political and economic institutions measured by $PIVdist$ and $EFWdist$ both negatively affect the probability of forming deep PTAs. Specifically, $EFWdist$ has a much stronger (negative) effect on the choice of deep PTAs than the simple formation of PTAs, indicating that homogeneity in economic institutions is crucial for trade partners to proceed to deep integration through the formation of comprehensive trade agreements that involve policy areas beyond trade such as competition policy and state-owned enterprises.

As both kinds of neighbour effect are statistically significant when included separately in the model, we next present the results based on the specification allowing for the mixed spillover effect in Table 8, so as to check whether the institutional spillovers are robust to the inclusion of geographical spillovers in the decisions of PTA type. Comparing the goodness of fit measured by negative log-likelihood, we find the optimal compound spatial weight matrix is obtained when α equals 0.1 as in column (1), which is further refined to 0.13 by the second round of grid search with finer grids. This set of results support the dominant role of institutional similarity in the spillovers in the type of PTAs, as the institutional channel accounts for around 90% in the mixed spillover effect.

The above findings are robust to the use of a panel data specification. As shown in column (2) of Table 9, the interdependence effect through institutional homogeneity is found to be an important determinant of the choice of PTA type. In column (3), we also confirm the existence of the spillovers generated by geographical proximity. Finally, in column (4), we find that the spillover effect of PTA type through institutional homogeneity is robust to the inclusion of that through geographical proximity as the estimate of $W^I_{t-5} \cdot PTA_{t-5}$ is still highly significant, with its value slightly decreasing from 1.963 to 1.594 after including $W^G_{t-5} \cdot PTA_{t-5}$ in the model.

4.2.2 Quantification

To have a better understanding of the spillovers in the decisions on the type of PTAs, we first present the marginal effects of the SARP estimation with the compound spatial weight matrix and compare the direct and the indirect ones. The related marginal effects are presented in the bottom half of Table 10. Consistent with the conclusion drawn from the estimated coefficients, the interdependence effect in PTA type is much stronger than that in PTA status, since the indirect effects for all the variables are more than double of the direct effects. For instance, the indirect effect of bilateral market size $GDPsum$ is 0.116, almost triple of the direct effect, suggesting that a 10 percent increase in the bilateral market size for a pair of trade partners increases the probability of having deep integration by around 1.16 percent for its institutional neighbours. Similarly, the difference in GDP per capita also has a more substantial indirect effect than direct effect on the choosing of deep PTAs: a 10 percent increase in the variable for a pair of trade partners leads to a 4.76% increase in the probability of forming deep PTAs for its neighbours sharing similar institutions.

Now we conduct two experiments by generating a one unit negative error shock to the latent regression model (1) for G7 and P5 country-pairs with respect to their relative utility gains from signing deep PTAs compared with signing shallow PTAs, separately, and we then examine its impact on the probabilities of forming deep PTAs among trade partners. In Figure 3, we plot the distributions of changes in probabilities of forming deep PTAs for G7 and non-G7 countries (under the G7 shock) and those for P5 and non-P5 countries (under the P5 shock), separately. First of all, we find that the changes in probabilities under both negative shocks have a wider range for the chosen PTA type as shown in Figure 3 than those for memberships as shown in Figure 2, where the largest drops in probabilities of forming deep PTAs are more than 70% and 60% under G7 shock and P5 shock, respectively. More importantly, the largest drops for non-G7 and non-G5 country-pairs are also significantly larger in this case, with the former greater than 40% and the latter greater than 30%, suggesting a stronger spillover effect of deep integration through institutional similarity from a quantitative perspective.[27] The mean values of the decreases in probabilities also support the above claim, which equal 53.5% and 13.9% for G7 and non-G7 country-pairs under the G7 shock, respectively, and 38.7% and 7.4% for

[27]Note that geographical proximity only accounts for 13% in the compound spatial weight matrix.

P5 and non-P5 country-pairs under the P5 shock, respectively. To sum up, the spillovers in the decisions on PTA type are overwhelming as relatively small countries are prone to the influences of large and institutionally similar countries' attitudes toward deep and shallow PTAs.

4.3 Robustness checks

4.3.1 Endogeneity issues

In the following, we analyze how robust the above results are with respect to the possible endogeneity in the spatial weights as well as independent variables, since there can be inverse causality between the formation of PTAs and domestic institutions. To be specific, on the one hand, countries are required to have institutional reform in a variety of policy fields, including protection of democratic rights, enforcement of intellectual property, regulation of labor market, and environmental standards, in order to be eligible to join certain trade agreements (e.g., the EU and NAFTA), and thus PTA memberships might affect institutional background. On the other hand, as investigated earlier, institutions can play an important role in the formation of PTAs.

As the existing methods do not allow for such endogeneity, we construct a new estimator based on the methods of Qu and Lee (2015) and Martinetti and Geniaux (2017). A small Monte Carlo simulation is conducted to assess its performance.[28] In the estimation, we use the ten-year lagged values of institutional variables as instruments, since ten years should be a long enough period for most trade agreement negotiation to complete. For instance, Wong and Yu (2015) find that the average duration for countries to join the WTO is around seven years with a standard deviation equal to five years.[29]

The estimation results are presented in Table 11 with the left panel of columns (1) and (2) addressing the first hypothesis about PTA memberships and the right panel of

[28]Please see the appendix for details about the new estimator and the Monte Carlo results.

[29]We also use other instruments such as five-year-lagged values of institutional variables, in which our results remain qualitatively unchanged.

columns (3) and (4) addressing the second hypothesis about the type of PTAs chosen. Columns (1) and (3) show the results given by the new estimation method that can deal with the potential endogeneity in both spatial weights and explanatory variables. It is worth noting that since we use the ten-year lagged values of democracy and economic freedom indexes as instrumental variables, we have a smaller sample size than the baseline one. Thus, the results presented by the new estimator might be driven by the sample size. To deal with this issue, we run the baseline SARP regressions with the same sample, as shown in columns (2) and (4). In general, our results suggest that endogeneity is not a big issue in shaping our findings as the estimates in columns (1) and (3) are very close to those in columns (2) and (4), respectively, in particular for the estimates of the spatial lag parameter ρ, confirming the significance of the institutional channel of spillovers in the formation of PTAs.

4.3.2 The EU member states

From the empirical point of view, another concern comes from the inclusion of all the EU member states separately in our regressions, which might compromise the generality of our findings as the EU member states account for a significant part of the PTA memberships in our sample. Thus, as another robustness check, we test the institutional spillover hypotheses by considering all the EU countries as a single entity.

As shown in Table 12, the key estimates related to the institutional interdependence remain statistically significant for either PTA memberships or their type. Still, there are a few minor changes in the results that are worth noting. First, in the formation of PTAs, the new estimates of the interdependence effects ρ are 0.215 and 0.318 for institutional and geographical channels, as shown in columns (1) and (2), respectively, which are relatively smaller than their counterparts in the baseline results that of 0.316 and 0.430 in Table 4, suggesting implicitly that the spillover effects of PTA memberships are relatively stronger among the EU countries than the rest of the world. Second, the results allowing for the mixed spillover effect as shown in column (3) indicate that institutional channel is less important in this case for PTA memberships, as the weight of geographical vicinity in the optimal compound spatial weight matrix α increases from 0.40 to 0.80. Finally, similar patterns are observed for the results of the type of PTAs in columns (4)-(6): the neighbour effects of PTA type are confirmed when we take all the EU countries as a single entity, although they are a little bit weaker in terms of magnitude; institutional similarity still dominates the mixed spillovers in the choice of the type of PTAs.

These new results not only confirm the robustness of our key findings regarding institutional spillovers in PTA formation, they are also consistent with the fact that PTAs reached by the EU countries tend to cover more non-trade issues than the rest of the world and thus can be more institutionally interdependent between each other (see, e.g., Horn et al. (2010)).

5 Conclusion

Our paper conducts an empirical examination of the relevance of institutions in the spillovers in PTA formation based on a formal spatial econometric model, and finds that the homogeneity of institutions such as democracy and economic freedom constitutes a new channel through which countries' decisions on PTA memberships and their type - deep or shallow agreement - can be interdependent.

Our findings highlight the significance of trade partners' institutional background in PTA formation in an interconnect world. More importantly, our findings are related to the debate on the role of PTAs in multilateralism since some economists see regionalism through forming PTAs as stumbling blocks while others see it as building blocks towards multilateralism. According to this study, the long-term effect of PTAs on global free trade, especially those deep agreements, to some extent, also depends on the global distribution of institutional classifications. The spillovers in trade agreement memberships, as initially suggested by Baldwin (1993), can either lead to multilateralism (or a two-pillar system as proposed in Baldwin (2016)) or several large trading blocs that share similar economic and political institutions within each bloc. We leave the empirical investigation of the welfare ramifications of our study for future research.

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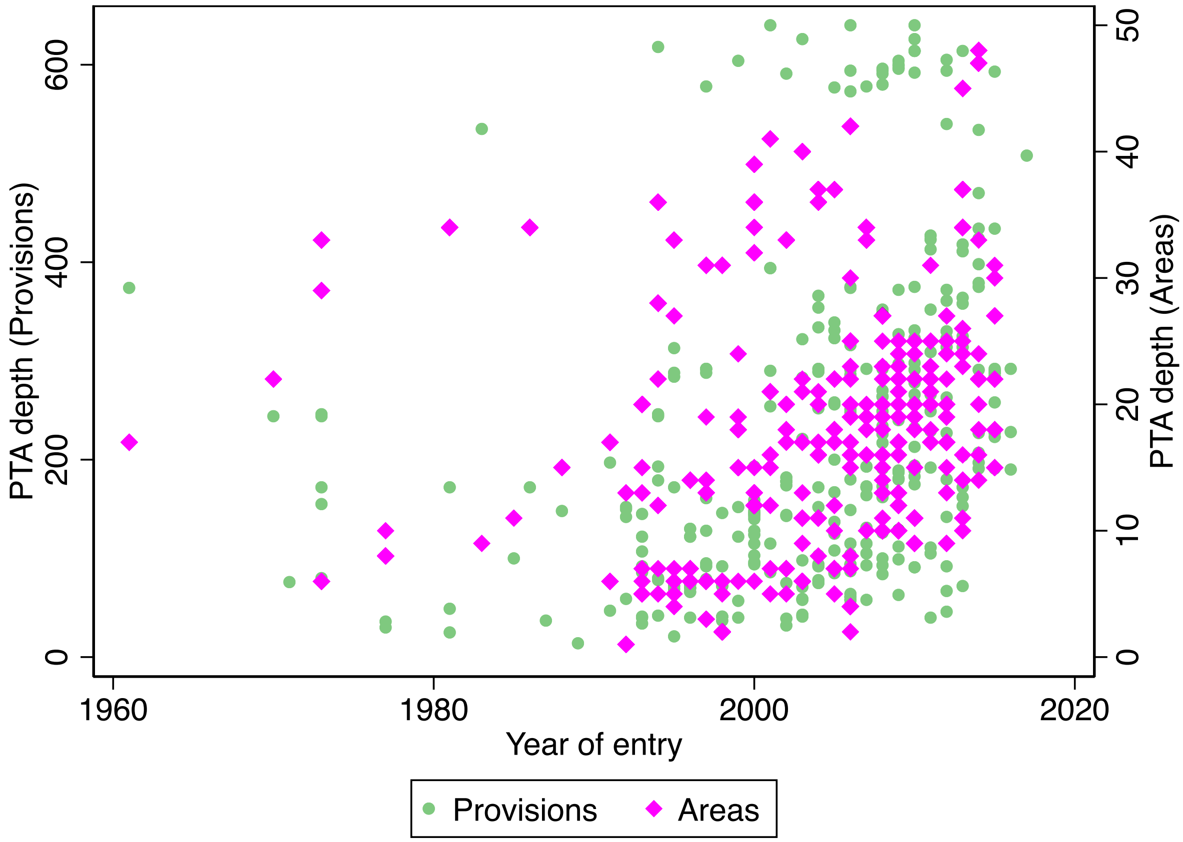
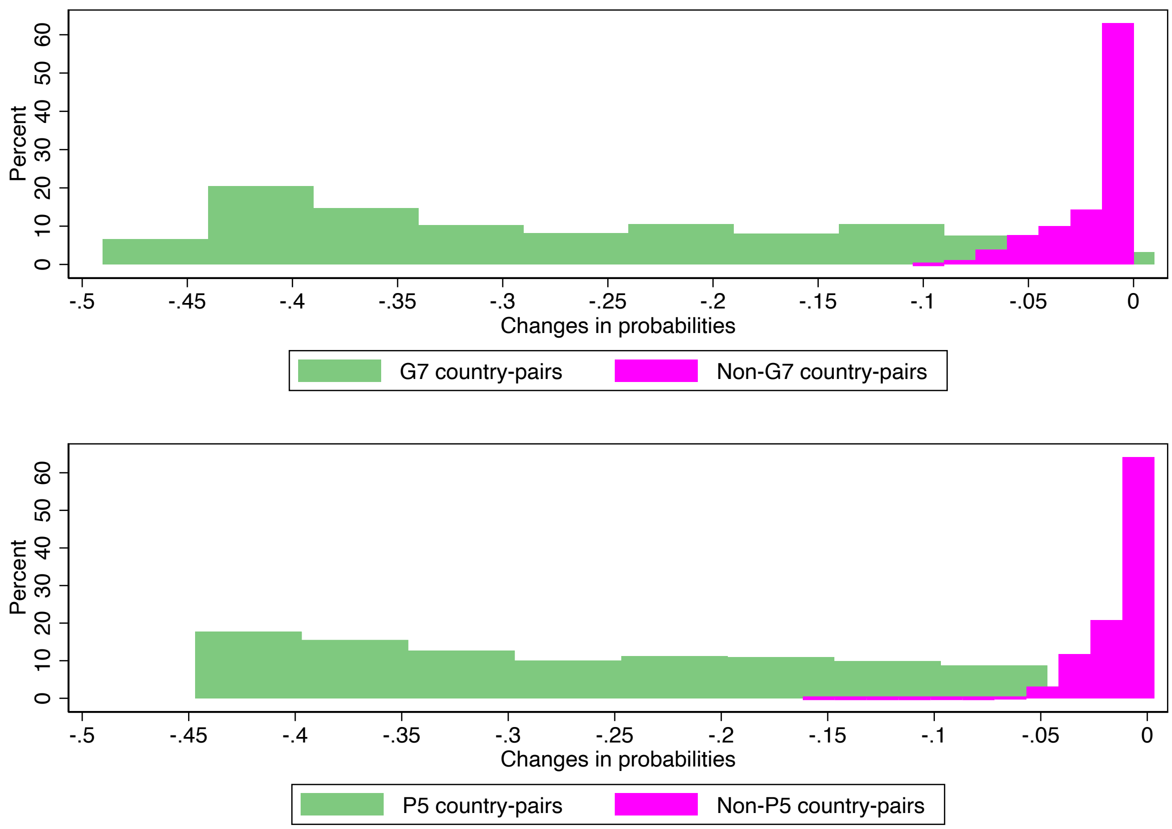


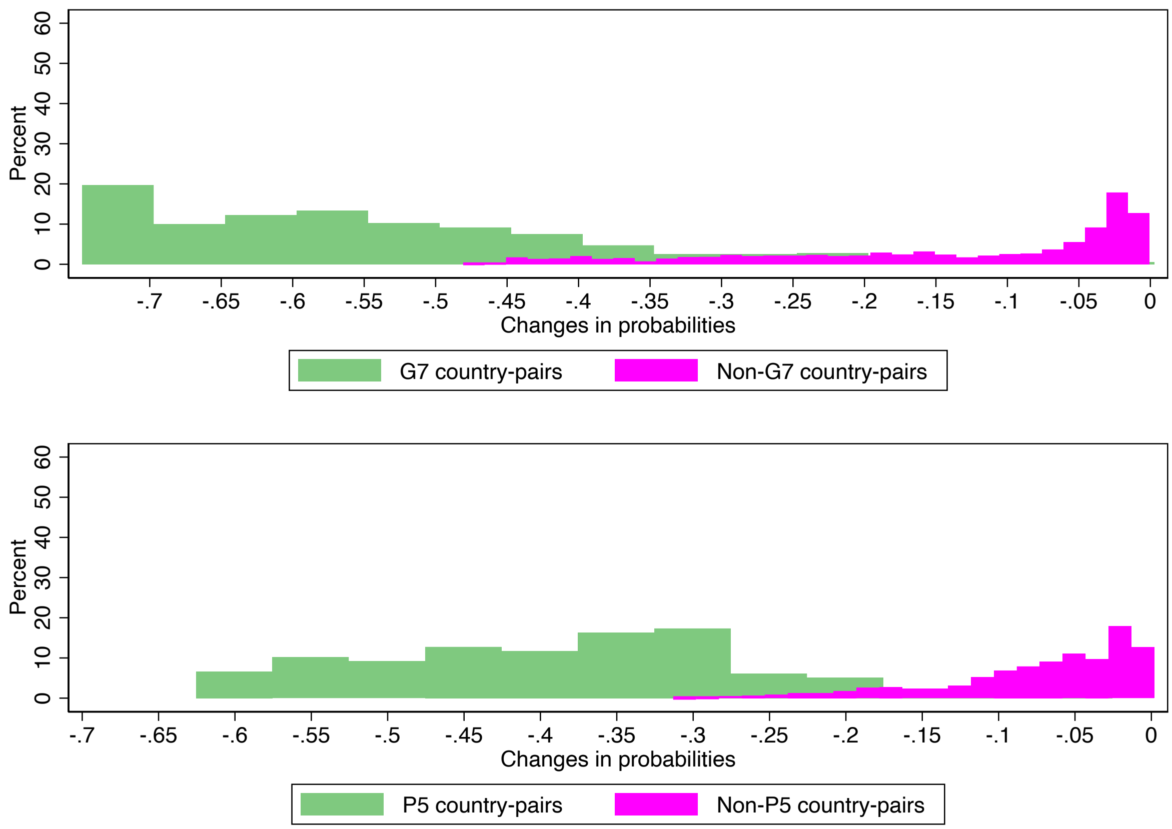
Note: This figure shows the evolution of PTA depth measured by both the number of provisions and areas between 1960 and 2017. The data of PTA areas is from Hofmann et al. (2017), and that of PTA provisions is from Mattoo et al. (2020).

Figure 1: The trend of PTA depth



Note: The top figure shows the distribution of probability changes in forming PTAs after a one unit negative error shock to the latent regression model (1) in forming PTAs among G7 countries and their trade partners. The top figure shows the distribution of probability changes in forming PTAs after a one unit negative error shock to the latent regression model (1) in forming PTAs among the five permanent members of the UN security council countries (P5) and their trade partners.

Figure 2: Quantification of deglobalization shocks to PTA memberships



Note: The top figure shows the distribution of probability changes in forming deep PTAs after a one unit negative error shock to the latent regression model (1) in forming PTAs among G7 countries and their trade partners, compared with shallow PTAs. The top figure shows the distribution of probability changes in forming PTAs after a one unit negative error shock to the latent regression model (1) in forming deep PTAs among the five permanent members of the UN security council countries (P5) and their trade partners.

Figure 3: Quantification of deglobalization shocks to the type of PTAs

Table 1: Sample correlations between PTA depth and institutional indicators

	Data source	PIVsum	EFWsum
Depth (# areas)	Hofmann et al. (2017)	0.543***	0.432***
Depth (# provisions)	Mattoo et al. (2020)	0.386***	0.536***

Note: This table presents the Pearson's correlations between PTA depth, measured by either the number of policy areas or provisions, and institutional scores. Specifically, PIVsum measures the summation of the trade partners' democracy scores, EFWsum measures the summation of the trade partners' economic freedom scores. Superscripts *, ** and *** represent statistical significance at the ten, five and one percent level, respectively.

Table 2: Determinants of PTA depth: The role of institutional indicators

	Depth (# areas) (1)	Depth (# provisions) (2)
NATURAL	-1.170***	-3.067
	(0.259)	(2.629)
GDPsum	-0.048	-8.214***
	(0.170)	(1.811)
GDPsim	0.595***	-2.508
	(0.224)	(2.337)
DKL	5.499***	44.508***
	(0.669)	(6.999)
SQDKL	-1.275***	-8.577***
	(0.176)	(1.891)
REMOTE	-30.441***	49.698***
	(1.836)	(18.729)
DROWKL	0.245	-12.454***
	(0.416)	(4.337)
PIVsum	0.630***	2.599***
	(0.032)	(0.329)
EFWsum	2.343***	47.926***
	(0.233)	(2.111)
Constant	245.222***	-814.354***
	(16.149)	(165.457)
R^2	0.435	0.333
Year	2017	2017
Observations	1,913	2,567

Note: See Table 4 for the details about the control variables. Standard errors are reported in parentheses. Superscripts *, ** and *** represent statistical significance at the ten, five and one percent level, respectively.

Table 3: Summary statistics

A. PTA memberships					
Variables	Observations (n)	Mean	Std. dev.	Min	Max
NATURAL	10,011	-8.674	0.783	-9.894	-4.088
GDPsum	10,011	26.324	1.679	21.583	31.085
GDPsim	10,011	-2.031	1.390	-9.165	-0.693
DKL	10,011	1.749	1.237	0.000	6.229
SQDKL	10,011	4.589	5.635	0.000	38.806
REMOTE	10,011	8.978	0.120	8.786	9.487
DROWKL	10,011	1.353	0.631	0.017	3.462
PIVdist	10,011	6.136	5.490	0.000	20.000
EFWdist	10,011	0.955	0.710	0.000	4.200
B. The type of PTAs					
Variables	Observations (n)	Mean	Std. dev.	Min	Max
NATURAL	2,956	-8.300	1.025	-9.892	-4.088
GDPsum	2,956	26.657	1.481	21.583	30.678
GDPsim	2,956	-1.812	1.167	-6.555	-0.693
DKL	2,956	1.519	1.077	0.000	5.544
SQDKL	2,956	3.466	4.400	0.000	30.735
REMOTE	2,956	8.960	0.131	8.786	9.487
DROWKL	2,956	1.196	0.601	0.023	3.462
PIVdist	2,956	4.285	4.609	0.000	19.000
EFWdist	2,956	0.841	0.668	0.000	4.200

Note: This table presents the summary statistics of the cross-sectional data in the year of 2017 that includes the variables of NATURAL (the log of the inverse of the great circle distance between two trade partners' capitals), REMOTE (the remoteness of a pair of continental trade partners from the rest of the world), GDPsum (total bilateral market size denotes the total GDP of trade partners), GDPsim (the similarity of two countries in terms of their GDP), DKL (the absolute difference in GDP per capita), SQDKL (the square of DKL), DROWKL (the relative factor endowment difference between the rest of the world and a given country-pair), PIVdist (the absolute distance of democracy between trade partners), and EFWdist (the absolute distance of economic freedom between trade partners).

Table 4: Empirical results: PTA memberships

	Theory (1)	Non-spatial (2)	W^I (3)	W^G (4)
$W^I \cdot PTA^*$	+		0.316***	
$W^G \cdot PTA^*$	+			0.430***
NATURAL	+	0.595***	0.583***	0.465***
		(0.023)	(0.023)	(0.018)
GDPsum	+	0.200***	0.179***	0.167***
		(0.011)	(0.011)	(0.009)
GDPsim	+	0.193***	0.179***	0.172***
		(0.014)	(0.013)	(0.012)
DKL	+	0.133***	0.134***	0.134***
		(0.040)	(0.039)	(0.038)
SQDKL	-	-0.025***	-0.029***	-0.033***
		(0.009)	(0.009)	(0.009)
REMOTE	+	0.653***	0.719***	0.421***
		(0.147)	(0.144)	(0.084)
DROWKL	-	-0.239***	-0.179***	-0.070***
		(0.029)	(0.027)	(0.021)
PIVdist	-	-0.060***	-0.044***	-0.055***
		(0.003)	(0.002)	(0.002)
EFWdist	-	-0.094***	-0.053***	-0.078***
		(0.022)	(0.020)	(0.020)
Negative Log-likelihood			4,949	4,687
Year		2017	2017	2017
Observations		10,011	10,011	10,011

Note: Column (1) presents the predicted signs from trade theory; columns (2)-(4) present the results with no spatial effect, the institutional spatial effect, and the geographical spatial effect, respectively. The control variables include NATURAL (the log of the inverse of the great circle distance between two trade partners' capitals), REMOTE (the remoteness of a pair of continental trade partners from the rest of the world), GDPsum (total bilateral market size denotes the total GDP of trade partners), GDPsim (the similarity of two countries in terms of their GDP), DKL (the absolute difference in GDP per capita), SQDKL (the square of DKL), DROWKL (the relative factor endowment difference between the rest of the world and a given country-pair), PIVdist (the absolute distance of democracy between trade partners), and EFWdist (the absolute distance of economic freedom between trade partners). Standard errors are reported in parentheses. Superscripts *, ** and *** represent statistical significance at the ten, five and one percent level, respectively.

Table 5: Empirical results with compound spatial weight matrix: PTA memberships

	$W^C = \alpha W^G + (1-\alpha) W^I$								
	$\alpha = 0.1$	$\alpha = 0.2$	$\alpha = 0.3$	$\alpha = 0.4$	$\alpha = 0.5$	$\alpha = 0.6$	$\alpha = 0.7$	$\alpha = 0.8$	$\alpha = 0.9$
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
$W^C \cdot PTA^*$	0.431***	0.438***	0.416***	0.410***	0.406***	0.403***	0.400***	0.396***	0.391***
NATURAL	0.463***	0.441***	0.451***	0.461***	0.464***	0.466***	0.464***	0.466***	0.462***
	(0.019)	(0.018)	(0.018)	(0.019)	(0.019)	(0.019)	(0.019)	(0.019)	(0.019)
GDPsum	0.179***	0.182***	0.179***	0.176***	0.177***	0.177***	0.177***	0.178***	0.178***
	(0.010)	(0.010)	(0.010)	(0.010)	(0.010)	(0.010)	(0.010)	(0.010)	(0.010)
GDPsim	0.178***	0.180***	0.178***	0.177***	0.178***	0.177***	0.177***	0.177***	0.177***
	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)
DKL	0.126***	0.124***	0.137***	0.143***	0.141***	0.142***	0.141***	0.138***	0.137***
	(0.039)	(0.039)	(0.039)	(0.039)	(0.039)	(0.039)	(0.039)	(0.039)	(0.039)
SQDKL	-0.034***	-0.037***	-0.038***	-0.038***	-0.037***	-0.037***	-0.037***	-0.036***	-0.035***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
REMOTE	0.270**	0.409***	0.531***	0.576***	0.593***	0.597***	0.601***	0.595***	0.585***
	(0.110)	(0.097)	(0.094)	(0.093)	(0.093)	(0.093)	(0.093)	(0.094)	(0.094)
DROWKL	-0.087***	-0.041*	-0.046**	-0.055**	-0.059**	-0.060**	-0.060**	-0.064***	-0.067***
	(0.024)	(0.023)	(0.023)	(0.023)	(0.023)	(0.023)	(0.023)	(0.023)	(0.024)
PIVdist	-0.039***	-0.044***	-0.049***	-0.052***	-0.054***	-0.055***	-0.056***	-0.057***	-0.057***
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.003)	(0.003)
EFWdist	-0.043**	-0.061***	-0.068***	-0.070***	-0.075***	-0.075***	-0.075***	-0.074***	-0.076***
	(0.020)	(0.020)	(0.020)	(0.021)	(0.021)	(0.021)	(0.021)	(0.021)	(0.021)
Negative Log-likelihood	4,854	4,802	4,804	4,800	4,802	4,801	4,805	4,808	4,815
Year	2017	2017	2017	2017	2017	2017	2017	2017	2017
Observations	10,011	10,011	10,011	10,011	10,011	10,011	10,011	10,011	10,011

Note: The table presents the results with compound spatial weight matrix W^C. The compound spatial weight matrix W^C is calculated from both the institutional spatial weight and the geographical spatial weight as shown in the appendix. See Table 4 for the details about the control variables. Standard errors are reported in parentheses. Superscripts *, ** and *** represent statistical significance at the ten, five and one percent level, respectively.

Table 6: Empirical results from the panel data: PTA memberships

	Non-spatial (1)	W^I (2)	W^G (3)	W^I & W^G (4)
$W^I_{t-5} \cdot PTA_{t-5}$		2.138***		1.854***
		(0.088)		(0.090)
$W^G_{t-5} \cdot PTA_{t-5}$			1.082***	1.010***
			(0.032)	(0.033)
NATURAL	0.690***	0.692***	0.588***	0.598***
	(0.015)	(0.015)	(0.015)	(0.015)
GDPsum	0.205***	0.159***	0.182***	0.144***
	(0.007)	(0.007)	(0.007)	(0.007)
GDPsim	0.197***	0.174***	0.179***	0.160***
	(0.009)	(0.009)	(0.009)	(0.010)
DKL	0.159***	0.151***	0.191***	0.182***
	(0.028)	(0.029)	(0.029)	(0.029)
SQDKL	-0.072***	-0.068***	-0.083***	-0.080***
	(0.007)	(0.007)	(0.007)	(0.007)
REMOTE	0.830***	0.964***	0.805***	0.915***
	(0.091)	(0.091)	(0.092)	(0.093)
DROWKL	-0.039**	-0.059***	-0.002	-0.022
	(0.017)	(0.018)	(0.018)	(0.018)
PIVdist	-0.024***	-0.008***	-0.022***	-0.009***
	(0.002)	(0.002)	(0.002)	(0.002)
EFWdist	-0.128***	-0.099***	-0.131***	-0.107***
	(0.013)	(0.013)	(0.013)	(0.013)
Pseudo R^2	0.222	0.244	0.263	0.278
Log-likelihood	-11104.243	-10802.488	-10524.768	-10311.113
Year	1996-2017	1996-2017	1996-2017	1996-2017
Observations	24,420	24,420	24,420	24,420

Note: The table presents the results obtained from the panel data specification. Columns (1)-(4) present the baseline results with no spatial effect, the institutional spatial effect, the geographical spatial effect, and both spatial effects, respectively. All the explanatory variables use the values observed in year $t-5$ in the panel data analysis with the details shown in the appendix. See Table 4 for the details about the control variables. Standard errors are reported in parentheses. Superscripts *, ** and *** represent statistical significance at the ten, five and one percent level, respectively.

Table 7: Empirical results: The type of PTAs

	Non-spatial (1)	W^I (2)	W^G (3)
$W^I \cdot PTA^*$		0.625***	
$W^G \cdot PTA^*$			0.571***
NATURAL	0.439***	0.393***	0.345***
	(0.031)	(0.029)	(0.023)
GDPsum	0.184***	0.126***	0.159***
	(0.020)	(0.017)	(0.013)
GDPsim	0.150***	0.113***	0.098***
	(0.026)	(0.024)	(0.022)
DKL	0.531***	0.557***	0.575***
	(0.081)	(0.078)	0.072
SQDKL	-0.102***	-0.107***	-0.112***
	(0.021)	(0.020)	(0.018)
REMOTE	-0.490**	-0.157	0.045
	(0.214)	(0.184)	(0.095)
DROWKL	-0.376***	-0.379***	-0.112***
	(0.049)	(0.040)	(0.029)
PIVdist	-0.042***	-0.015***	-0.049***
	(0.006)	(0.003)	(0.005)
EFWdist	-0.258***	-0.170***	-0.233***
	(0.042)	(0.033)	(0.037)
Negative Log-likelihood		1,489	1,439
Year	2017	2017	2017
Observations	2,956	2,956	2,956

Note: Columns (1)-(3) present the baseline results with no spatial effect, the institutional spatial effect, and the geographical spatial effect, respectively. See Table 4 for the details about the control variables. Superscripts *, ** and *** represent statistical significance at the ten, five and one percent level, respectively.

Table 8: Empirical results with compound spatial weight matrix: The type of PTAs

	$W^C = \alpha W^G + (1-\alpha)W^I$								
	$\alpha = 0.1$	$\alpha = 0.2$	$\alpha = 0.3$	$\alpha = 0.4$	$\alpha = 0.5$	$\alpha = 0.6$	$\alpha = 0.7$	$\alpha = 0.8$	$\alpha = 0.9$
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
$W^C \cdot PTA^*$	0.797***	0.815***	0.791***	0.788***	0.763***	0.743***	0.726***	0.721***	0.708***
NATURAL	0.364***	0.342***	0.343***	0.351***	0.357***	0.354***	0.354***	0.356***	0.353***
	(0.022)	(0.021)	(0.022)	(0.022)	(0.023)	(0.023)	(0.023)	(0.023)	(0.023)
GDPsum	0.142***	0.162***	0.176***	0.184***	0.191***	0.192***	0.194***	0.193***	0.191***
	(0.014)	(0.013)	(0.012)	(0.012)	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)
GDPsim	0.022	0.074***	0.120***	0.148***	0.156***	0.156***	0.155***	0.154***	0.152***
	(0.021)	(0.022)	(0.023)	(0.023)	(0.024)	(0.024)	(0.024)	(0.024)	(0.024)
DKL	0.621***	0.608***	0.594***	0.577***	0.579***	0.586***	0.583***	0.580***	0.578***
	(0.075)	(0.077)	(0.078)	(0.079)	(0.080)	(0.080)	(0.080)	(0.081)	(0.081)
SQDKL	-0.136***	-0.136***	-0.129***	-0.122***	-0.123***	-0.126***	-0.126***	-0.125***	-0.125***
	(0.018)	(0.019)	(0.019)	(0.020)	(0.020)	(0.020)	(0.021)	(0.021)	(0.021)
REMOTE	0.172*	0.289***	0.400***	0.523***	0.572***	0.578***	0.568***	0.571***	0.549***
	(0.092)	(0.077)	(0.074)	(0.072)	(0.075)	(0.077)	(0.079)	(0.080)	(0.082)
DROWKL	-0.111***	-0.033	0.013	0.055**	0.066***	0.065**	0.061**	0.057**	0.046*
	(0.030)	(0.027)	(0.026)	(0.025)	(0.025)	(0.026)	(0.026)	(0.027)	(0.027)
PIVdist	-0.004	-0.009***	-0.020***	-0.029***	-0.036***	-0.039***	-0.041***	-0.042***	-0.043***
	(0.003)	(0.003)	(0.003)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
EFWdist	-0.063*	-0.093***	-0.160***	-0.219***	-0.241***	-0.250***	-0.253***	-0.254***	-0.251***
	(0.034)	(0.034)	(0.035)	(0.037)	(0.037)	(0.038)	(0.038)	(0.038)	(0.039)
Negative Log-likelihood	1319	1335	1352	1362	1373	1385	1392	1397	1410
Year	2017	2017	2017	2017	2017	2017	2017	2017	2017
Observations	2,956	2,956	2,956	2,956	2,956	2,956	2,956	2,956	2,956

Note: The table presents the results with compound spatial weight matrix W^C. The compound spatial weight matrix W^C is calculated from both the institutional spatial weight and the geographical spatial weight as shown in the appendix. See Table 4 for the details about the control variables. Standard errors are reported in parentheses. Superscripts *, ** and *** represent statistical significance at the ten, five and one percent level, respectively.

Table 9: Empirical results from the panel data: The type of PTAs

	Non-spatial (1)	W^I (2)	W^G (3)	W^I & W^G (4)
$W^I_{t-5} \cdot PTA_{t-5}$		1.963***		1.594***
		(0.095)		(0.101)
$W^G_{t-5} \cdot PTA_{t-5}$			1.260***	0.929***
			(0.067)	(0.072)
NATURAL	0.756***	0.730***	0.677***	0.669***
	(0.032)	(0.033)	(0.033)	(0.034)
GDPsum	0.127***	-0.000	0.104***	0.005
	(0.019)	(0.021)	(0.020)	(0.022)
GDPsim	0.203***	0.116***	0.149***	0.087***
	(0.028)	(0.030)	(0.029)	(0.031)
DKL	-0.700***	-0.623***	-0.742***	-0.672***
	(0.093)	(0.098)	(0.098)	(0.100)
SQDKL	0.257***	0.217***	0.267***	0.235***
	(0.029)	(0.031)	(0.031)	(0.031)
REMOTE	-1.149***	-0.355*	-1.539***	-0.825***
	(0.186)	(0.193)	(0.197)	(0.202)
DROWKL	-0.166***	-0.212***	-0.149***	-0.206***
	(0.044)	(0.046)	(0.046)	(0.048)
PIVdist	0.011*	0.036***	0.017***	0.035***
	(0.006)	(0.007)	(0.006)	(0.007)
EFWdist	-0.383***	-0.264***	-0.377***	-0.288***
	(0.037)	(0.039)	(0.038)	(0.040)
Pseudo R^2	0.346	0.432	0.415	0.464
Log-likelihood	-1744.566	-1514.032	-1559.878	-1430.697
Year	1996-2017	1996-2017	1996-2017	1996-2017
Observations	3,912	3,912	3,912	3,912

Note: The table presents the results obtained from the panel data specification. Columns (1)-(4) present the baseline results with no spatial effect, the institutional spatial effect, the geographical spatial effect, and both spatial effects, respectively. All the explanatory variables use the values observed in year $t-5$ in the panel data analysis with the details shown in the appendix. See Table 4 for the details about the control variables. Standard errors are reported in parentheses. Superscripts *, ** and *** represent statistical significance at the ten, five and one percent level, respectively.

Table 10: Marginal effects

A. PTA memberships

Variables	Direct	Indirect	Total
NATURAL	0.129	0.088	0.217
GDPsum	0.049	0.034	0.083
GDPsim	0.049	0.034	0.083
DKL	0.040	0.027	0.067
SQDKL	-0.011	-0.007	-0.018
REMOTE	0.161	0.110	0.271
DROWKL	-0.015	-0.011	-0.026
PIVdist	-0.015	-0.010	-0.025
EFWdist	-0.020	-0.013	-0.033

B. The type of PTAs

Variables	Direct	Indirect	Total
NATURAL	0.100	0.276	0.376
GDPsum	0.042	0.116	0.158
GDPsim	0.009	0.024	0.032
DKL	0.173	0.476	0.649
SQDKL	-0.039	-0.106	-0.145
REMOTE	0.061	0.168	0.229
DROWKL	-0.023	-0.064	-0.087
PIVdist	-0.001	-0.003	-0.004
EFWdist	-0.019	-0.053	-0.073

Note: The table presents the marginal effects obtained from the SARP model allowing for the mixed spillover effect. Direct effects measures the average over all the observations of the **own** marginal effect of a control variable on the probability for PTA formation. Indirect effects represent the average over all the observations of the marginal effect of an explanatory variable on the choice probability of all other observations. Total effects stand for the sum of direct and indirect impacts. See Table 4 for the details about the control variables.

Table 11: Robustness checks: Endogeneity

	PTA memberships		The type of PTAs	
	Endogeneity	Baseline	Endogeneity	Baseline
	(1)	(2)	(3)	(4)
$W^I \cdot PTA^*$	0.303***	0.301***	0.621***	0.629***
NATURAL	0.587***	0.573***	0.445***	0.392***
	(0.026)	(0.026)	(0.032)	(0.030)
GDPsum	0.169***	0.161***	0.112***	0.109***
	(0.012)	(0.012)	(0.019)	(0.017)
GDPsim	0.178***	0.174***	0.112***	0.112***
	(0.016)	(0.016)	(0.025)	(0.024)
DKL	0.114***	0.108**	0.669***	0.600***
	(0.044)	(0.043)	(0.083)	(0.081)
SQDKL	-0.013	-0.016	-0.140***	-0.126***
	(0.010)	(0.010)	(0.021)	(0.021)
REMOTE	0.453***	0.454***	-0.110	-0.320*
	(0.160)	(0.158)	(0.209)	(0.191)
DROWKL	-0.247***	-0.230***	-0.382***	-0.367***
	(0.030)	(0.029)	(0.044)	(0.040)
PIVdist	-0.048***	-0.047***	-0.022***	-0.017***
	(0.004)	(0.002)	(0.005)	(0.003)
EFWdist	-0.183***	-0.090***	-0.227***	-0.119***
	(0.036)	(0.024)	(0.046)	(0.036)
Negative log-likelihood	4023	4068	1379	1390
Year	2017	2017	2017	2017
Observations	7,875	7,875	2,772	2,772

Note: The table presents the results obtained from the SARP model that allows for the endogeneity in both spatial weights and explanatory variables. See Table 4 for the details about the control variables. Standard errors are reported in parentheses. Superscripts *, ** and *** represent statistical significance at the ten, five and one percent level, respectively.

Table 12: Robustness checks: The EU countries as a single entity

	PTA memberships			The type of PTAs		
	W^I	W^G	$W^C(\alpha = 0.8)$	W^I	W^G	$W^C(\alpha = 0.2)$
	(1)	(2)	(3)	(4)	(5)	(6)
$W^I \cdot PTA^*$	0.215***			0.584***		
$W^G \cdot PTA^*$		0.318***			0.496***	
$W^C \cdot PTA^*$			0.267***			0.753***
NATURAL	0.721***	0.597***	0.605***	0.416***	0.417***	0.340***
	(0.028)	(0.023)	(0.024)	(0.036)	(0.034)	(0.031)
GDPsum	0.303***	0.271***	0.287***	0.245***	0.188***	0.206***
	(0.015)	(0.013)	(0.014)	(0.025)	(0.022)	(0.020)
GDPsim	0.264***	0.243***	0.255***	0.222***	0.146***	0.206***
	(0.018)	(0.017)	(0.017)	(0.035)	(0.033)	(0.034)
DKL	0.130**	0.104*	0.112**	0.410***	0.355***	0.509***
	(0.054)	(0.053)	(0.054)	(0.112)	(0.111)	(0.109)
SQDKL	-0.067***	-0.061***	-0.062***	-0.049	-0.031	-0.066**
	(0.014)	(0.014)	(0.014)	(0.031)	(0.030)	(0.029)
REMOTE	1.995***	1.181***	1.418***	1.778***	1.186***	0.803***
	(0.170)	(0.120)	(0.131)	(0.240)	(0.143)	(0.121)
DROWKL	0.139***	0.138***	0.146***	0.099*	0.121***	0.122***
	(0.033)	(0.028)	(0.030)	(0.055)	(0.043)	(0.036)
PIVdist	-0.029***	-0.032***	-0.035***	0.001	-0.006	-0.010**
	(0.003)	(0.003)	(0.003)	(0.004)	(0.006)	(0.005)
EFWdist	0.057**	0.056**	0.051*	-0.167***	-0.122**	-0.166***
	(0.027)	(0.026)	(0.026)	(0.050)	(0.051)	(0.050)
Negative log-likelihood	2971	2903	2952	765	759	707
Year	2017	2017	2017	2017	2017	2017
Observations	7,118	7,118	7,118	1,551	1,551	1,551

Note: The table presents the results obtained from the SARP model that considers all the EU member countries as a single entity. See Table 4 for the details about the control variables. Standard errors are reported in parentheses. Superscripts *, ** and *** represent statistical significance at the ten, five and one percent level, respectively.

Appendix A. Econometric Issues and Country Coverage (Not for Publication)

Compound spatial weight matrix

We revise the SARP model, in particular the spatial weight matrix, to allow for the mixed domino effect generated from both institutional and geographical proximities in the empirical analysis. Thus, we construct a compound spatial weight matrix as follows by following Case et al. (1993) and the recent literature in real estate such as Zhang et al. (2019):

$$W^C = \alpha W^G + (1-\alpha)W^I, \alpha \in [0,1], \tag{A.1}$$

where W^I is the spatial weight matrix constructed from institutional components and W^G is the spatial weight matrix constructed from geographical components. We define α as the proportion of spatial weight constructed from geographic distance, ranging from 0 to 1. Thus, the contribution of institutional differential to the compound spatial weighting decreases with α, while that of geographical distance increases with α. In order to calculate the spatial weights, we need to calculate the geographical distance between, for instance, country-pairs i and j as follows:

$$Distance^G_{ij} = (\sum_{c_i}\sum_{c_j} Distance^G_{c_i,c_j})/4 \tag{A.2}$$

where $c_i \in \{i_1, i_2\}$ denotes a country from the country-pair i and $c_j \in \{j_1, j_2\}$ denotes a country from the country-pair j, respectively. The term $Distance^G_{c_i,c_j}$ measure the geographical distance between countries c_i and c_j, which is calculated from the latitude and longitude coordinates of their capitals. Thus, the term $\sum_{c_i}\sum_{c_j} Distance^G_{c_i,c_j}$ measures the total geographical distance between the members of the two country-pairs i and j, and it is further divided by four to obtain the average distance between the two country-pairs.

Based upon the average geographical distance between country-pairs, a typical component of W^G without row-standardization is calculated from a negative exponential function :

$$w_{i,j}^{G*} = e^{-Distance_{ij}^{G}/500}, i \neq j, \tag{A.3}$$

which denotes the $(i,j)^{th}$ element in the negative exponential distance weight matrix, with all the diagonal entries set to zeros. Again, we replace the cells of the spatial weight matrix that are less than the value of the 95th percentile with zeros to save memory space in estimation. Next, we make row-standarization for the above matrix to obtain the actual elements of the geographical spatial weight matrix W^G:

$$w_{i,j}^{G} = w_{i,j}^{G*} / \sum_{j} w_{i,j}^{G*}, \tag{A.4}$$

which denotes the $(i,j)^{th}$ element in the final spatial weight matrix W^G, with the sum of each row equal to one.

Panel data specification

For the benefits of robustness, we also use the standard probit model with a dynamic spatial lag term that of a time varying spatial weight matrix as one of the regressors to test the institutional interdependence hypothesis in the panel data. In this case, we use data on membership events for four years between 1996 and 2017 in a manner that $t = 2001, 2006, 2011, 2017$, and $t - 5 = 1996, 2001, 2006, 2012$. The intervals are five years but with one exception to make full use of the sample. Table A.1 shows the summary statistics for the panel data we use. Adopting the common assumption in the literature that new memberships of PTAs in the present do not affect the probability of new memberships in the past such as five years ago, we have the panel probit model in vector form as follows:

$$\begin{aligned} PTA^*{}_t &= \rho W^I_{t-5} \cdot PTA_{t-5} + X_{t-5}\beta + \varepsilon_t, \\ PTA_{t-5} &= I[PTA^*{}_{t-5} > 0], \end{aligned} \tag{A.5}$$

where $PTA^*{}_t$ and PTA_{t-5} are $n \times 1$ vectors, 0 is an $n \times 1$ vector of zeros, W^I_{t-5} is an $n \times n$ spatial weight matrix constructed from institutional differential observed five years ago, X_{t-5} is an $n \times k$ matrix of other explanatory variables, ρ is an unknown parameter by measuring the spatial effect of institutional homogeneity, and β is a vector of unknown parameters of other explanatory variables. Finally ε_t is an error term and n stands for the number of country-pairs (or observations in each year). As before, PTA_t is a binary vector which takes value 1 if there is a preferential trade agreement within a country-pair in the year of t and takes value 0 otherwise, and $PTA^*{}_t$ is still an unobservable vector that measures the minimum welfare gains of a pair of trade partners for signing a PTA between each other. Again, X_{t-5} includes all the control variables such as $NATURAL$, $REMOTE$, $GDPsum$, $GDPsim$, DKL, $SQDKL$, $DROWKL$, $PIVdist$, $EFWdist$, and a constant term. As defined in the cross-sectional case, the entries of the spatial weight matrix are inversely related to the institutional differential calculated from the indicators of democracy and economic freedom, with all the diagonal elements set to zeros. Particularly in the case of panel data analysis, we calculate the spatial lag term $W^I_{t-5} \cdot PTA_{t-5}$ beforehand and treat it as one of the regressors to account for the spatial interdependence effect in the probit model. The advantages of having a panel data model include simpler estimation and larger size of data. In addition, the regression results drawn from the panel data model in Equation (A.5) have more implications about the dynamic feature of the evolution of PTAs than the cross-sectional ones which mainly estimate the long-term effect. In testing the second hypothesis about the type of PTAs chosen, we define PTA_{t-5} as the status of deep or shallow PTAs instead, which equals 1 if there is a deep PTA and 0 if there is a shallow PTA.

Inference of spatial autoregressive probit model

In this section, we introduce the inference of the SARP model we follow to calculate the marginal effects and spillovers.

Consider a SARP model

$$\begin{aligned} PTA^* &= \rho W^I \cdot PTA^* + X\beta + \varepsilon, \\ PTA &= I[PTA^* > 0]. \end{aligned} \tag{A.6}$$

It then follows

$$\begin{aligned} PTA^* &= [I_n - \rho W]^{-1}[X\beta + \varepsilon], \\ PTA_i^* &= e_i'[I_n - \rho W]^{-1}[X\beta + \varepsilon], \end{aligned} \tag{A.7}$$

where e_i is the ith column vector of the identity matrix I_n, and

$$\begin{aligned} Pr(PTA_i = 1|X) &= Pr(e_i'[I_n - \rho W]^{-1}[X\beta + \varepsilon] > 0) \\ &= Pr(e_i'u > -e_i'[I_n - \rho W]^{-1}X\beta|X) \\ &= \Phi_i(e_i'[I_n - \rho W]^{-1}X\beta), \end{aligned} \tag{A.8}$$

where $\varepsilon \sim N(0_n, I_n)$ and $u = [I_n - \rho W]^{-1}\varepsilon \sim N(0_n, \Sigma_n)$ with $\Sigma_n = [I_n - \rho W]^{-1}[I_n - \rho W']^{-1}$, and $e_i'u_i \sim N(0, \sigma_i^2)$ with $\sigma_i^2 = e_i'\Sigma_n e_i$ and $\Phi_i(\cdot)$ is the cumulative distribution function of a normal random variable with zero mean and variance σ_i^2.

Then, the marginal effect of kth regressor is given by

$$\begin{aligned} \frac{\partial Pr(PTA_i = 1|X)}{\partial x_k'} &= \frac{\partial \Phi_i(e_i'[I_n - \rho W]^{-1}X\beta)}{\partial x_k'} \\ &= \phi_i(e_i'[I_n - \rho W]^{-1}X\beta)[I_n - \rho W']^{-1}e_i\beta_k, \end{aligned} \tag{A.9}$$

where $x_k = [x_{1k}, ..., x_{nk}]'$ is an $n \times 1$ vector, and $\phi_i(\cdot)$ is the density function of a normal random variable with zero mean and variance σ_i^2. Denoting $\alpha_i = \phi_i(e_i'[I_n -]^{-1}X\beta)$, we therefore have

$$\frac{\partial Pr(PTA=1|X)}{\partial x_k'} = \begin{bmatrix} \alpha_1 e_1'[I_n-\rho W]^{-1} \\ \vdots \\ \alpha_n e_n'[I_n-\rho W]^{-1} \end{bmatrix} \beta_k$$
$$= A_n[I_n-\rho W]^{-1}\beta_k$$
$$= \Delta_{n,k} \quad \text{(A.10)}$$

where $A_n = diag(\alpha_1, \cdots, \alpha_n)$, and we have

1. Direct impact: the diagonal elements in matrix $\Delta_{n,k}$.

2. Indirect impacts to country-pair i: the ith row of $\Delta_{n,k}$ except for the $(i;i)$th element of $\Delta_{n,k}$. This indirect impact considers how changes of kth explanatory variable in all country-pairs influence country-pair i.

3. Indirect impact from country-pair i: the ith column of $\Delta_{n,k}$ except for the $(i;i)$th element of $\Delta_{n,k}$. This form of indirect impacts measures how changes of the kth explanatory variable in country-pair i influence all other country-pairs.

Note, the average total of these two forms of indirect impacts are numerically equal, although they represent different interpretative viewpoints.

Next, we consider the impact of a unit shock from error term ε_i to the dependent variable. Let $\hat{F}_i = \Phi_i(e_i'[I_n-\hat{\rho}W]^{-1}X\hat{\beta})$ be the estimate of $F_i = Pr(PTA_i=1|X) = \Phi_i(e_i'[I_n-\rho W]^{-1}X\beta)$. Then, a popularly used method sets the fitted dependent variable $P\hat{T}A_i = 1$ if $\hat{F}_i > 0.5$ and $P\hat{T}A_i = 0$ otherwise. Consider that a shock (c) to the error term ε, which instantly yields the change of expected latent utility of

$$\Delta PTA^* = [I_n-\rho W]^{-1}c. \quad \text{(A.11)}$$

Then, after the shock, the probability of $PTA_i = 1$ becomes

$$\begin{aligned} F_{i,new} &= Pr(PTA_i^* + \Delta PTA_i^* > 0|X) \\ &= Pr(e_i'[I_n - \rho W]^{-1}[X\beta + \varepsilon] > -\Delta PTA_i^*) \\ &= Pr(e_i'u > -e_i'[I_n - \rho W]^{-1}X\beta - \Delta PTA_i^*|X) \\ &= \Phi_i(e_i'[I_n - \rho W]^{-1}(X\beta + c), \end{aligned} \tag{A.12}$$

which is to estimated by $\hat{F}_{i,new} = \Phi_i(e_i'[I_n - \hat{\rho}W]^{-1}(X\hat{\beta} + c))$.

An estimation procedure allowing for endogeneity

Following Qu and Lee (2015), we set up a new estimation procedure to deal with the endogeneity issues in the following model in matrix form:

$$\begin{aligned} &PTA^* = \rho W^I \cdot PTA^* + X_1\beta + \varepsilon, \\ &PTA = I[PTA^* > 0], \end{aligned} \tag{A.13}$$

where the notation is generally consistent with that in Section 3. Specifically, $X_1 = [X_{11}, X_{12}]$, where X_{12} is endogenous that contains the institutional control variables (i.e., PIVdist, EFWdist) and X_{11} is exogenous that contains other control variables. The institutional spatial weight matrix W^I has its typical element $w_{ij}^I = h(z_{ij})$, where z_{ij} is the institutional differential between country-pairs calculated from the country-pair aggregate scores of democracy and economic freedom, following Equation (2), and h represents the negative exponential function. We let Z denote the two aggregate indicators of democracy and economic freedom that generate z_{ij} and we further assume that Z is endogeneous. Thus, the spatial weight matrix W^I is allowed to be endogeneous in our model because of the endogeneity of Z. Next, we generate an auxiliary model as follows:

$$X_{12} = X_2\Gamma_1 + u_1, \tag{A.14}$$

and

$$Z = X_3\Gamma_2 + u_2, \tag{A.15}$$

where X_2 and X_3 are exogenous variables, and X_{11}, X_2 and X_3 are allowed to share common variables. In other words, X_{11}, X_2, and X_3 are exogenous to u_1, u_2, and ε. Next, we make some regularity assumption about the distributions of the error terms u_1, u_2, and ε as follows:

$$\begin{pmatrix} u_{1i} \\ u_{2i} \\ \varepsilon_i \end{pmatrix} \overset{i.i.d.}{\sim} N(0, \Sigma), \tag{A.16}$$

where

$$\Sigma = \begin{pmatrix} \Sigma_{u_1}^2 & \Sigma'_{u_1 u_2} & \Sigma'_{\varepsilon u_1} \\ \Sigma_{u_1 u_2} & \Sigma_{u_2}^2 & \Sigma'_{\varepsilon u_2} \\ \Sigma_{\varepsilon u_1} & \Sigma_{\varepsilon u_2} & \sigma_\varepsilon^2 \end{pmatrix}. \tag{A.17}$$

Next, letting $\xi = \varepsilon - u_1\delta_1 - u_2\delta_2$, $\delta_1 = \Sigma^{-1}\Sigma_{\varepsilon u_1}$, and $\delta_2 = \Sigma^{-1}\Sigma_{\varepsilon u_2}$, we have:

$$PTA^* = \rho W^I \cdot PTA^* + X_1\beta + (X_{12} - X_2\Gamma_1)\delta_1 + (Z - X_3\Gamma_2)\delta_2 + \xi, \tag{A.18}$$

where $\sigma_\xi^2 = \sigma_\varepsilon^2 - \Sigma'_{\varepsilon u_1}\Sigma_{u1}^{2-1}\Sigma_{\varepsilon u_1} - \Sigma'_{\varepsilon u_2}\Sigma_{u2}^{2-1}\Sigma_{\varepsilon u_2}$.

After running the regressions based on Equations (A.14) and (A.15) (or first-step regressions), we obtain the estimates of Γ_1 and Γ_2 which are denoted by $\widehat{\Gamma}_1$ and $\widehat{\Gamma}_2$, respectively. Next, we replace Γ_1 and Γ_2 by $\widehat{\Gamma}_1$ and $\widehat{\Gamma}_2$, respectively, in Equation (A.18), and estimate the SARP model by approximate maximum likelihood estimation method in the second-step regression (Martinetti and Geniaux, 2017). A set of simulations related to the estimation method is presented in the following section.

We present the main results of the Monte Carlo simulations that evaluate the finite sample performance of the estimation procedure introduced above that allows for the endogeneity in both the spatial weights and some explanatory variables. In the experiment, the sample size n varies among 100, 250, 500, and 1,000 to measure the performance of the method in terms of estimation consistency. For each n, we repeat 1,000 Monte Carlo trials. The true values in the data generation process are set as $\beta_0 = 0$, $\beta_1 = 0.3$, $\beta_2 = 0.8$, $\Gamma_1 = 1$, $\Gamma_2 = 1$, and $\rho = 0.3$, without loss of generality. Endogeneity is induced in the model by assuming that ε is correlated with u_1 and u_2. Specifically, we assume $\varepsilon = u_1 + u_2 + u_3$, where u_1, u_2, and u_3 follow the standard normal distribution independently. The spatial weight matrix is calculated as the inverse Euclidean distance of the endogeneous variable Z, with row-standardization and all the diagonals and the entries that are less than the value of the 90th percentile set as zeros. One notes that the simulation results are qualitatively unchanged if we choose other threshold percentiles such as the 95th percentile. The estimated results given in Table A.2 support the consistency of the proposed estimator. The Root Mean Square Error (RMSE) drops quickly toward zero for all the four parameters of interest $(\beta_0, \beta_1, \beta_2, \rho)$ as the sample size increases. In particular, the RMSE for the estimate of the interdependence effect ρ decreases from 0.113 to 0.048 when sample sizes increases from 100 to 1,000. Overall, the results from the above simulations imply that the estimator given by the constructed two-step estimation method is consistent in the presence of endogeneity in both the spatial weight matrix and explanatory variables. In another experiment, we show that the estimator we use in the baseline regressions is not consistent if there is endogeneity in either the spatial weight matrix or explanatory variables, whose results are available upon request.

Table A.1: Summary statistics for the panel data

A. PTA memberships

Variables	N	Mean	Std. dev.	Min	Max
NATURAL	24,420	-8.698	0.814	-9.892	-4.088
GDPsum	24,420	25.991	1.771	20.699	30.838
GDPsim	24,420	-2.085	1.427	-9.614	-0.693
DKL	24,420	1.815	1.280	0.000	6.152
SQDKL	24,420	4.935	5.952	0.000	37.850
REMOTE	24,420	8.988	0.128	8.786	9.487
DROWKL	24,420	1.440	0.664	0.013	3.640
PIVdist	24,420	5.785	5.312	0.000	20.000
EFWdist	24,420	1.143	0.886	0.000	5.770

B. PTA types

Variables	N	Mean	Std. dev.	Min	Max
NATURAL	3,912	-7.969	1.117	-9.891	-4.088
GDPsum	3,912	26.306	1.601	20.916	30.520
GDPsim	3,912	-1.646	1.024	-6.403	-0.693
DKL	3,912	1.130	0.858	0.001	4.221
SQDKL	3,912	2.014	2.722	0.000	17.816
REMOTE	3,912	8.964	0.158	8.788	9.487
DROWKL	3,912	1.410	0.652	0.028	3.640
PIVdist	3,912	4.081	4.746	0.000	17.000
EFWdist	3,912	0.915	0.796	0.000	5.770

Note: This table presents the summary statistics of the panel data over 1996 to 2017. We have 6,105 observations for each year and totally 24,420 observations for four years. See Table 3 for the detailed information about the variables.

Table A.2: Simulation results for the estimation method with endogeneity

	β_0	β_1	β_2	ρ
A. $n = 100$				
Avg. est.	0.059964	0.396611	1.124607	0.281456
Avg. bias	0.059964	0.096611	0.324607	-0.01854
Avg. abs. bias	0.301337	0.266906	0.39073	0.088301
Std. dev.	2.01813	0.421497	3.039674	0.111779
RMSE	2.018012	0.432222	3.055446	0.113252
B. $n = 250$	β_0	β_1	β_2	ρ
Avg. est.	0.003735	0.31652	0.852911	0.295078
Avg. bias	0.003735	0.01652	0.052911	-0.00492
Avg. abs. bias	0.117819	0.118716	0.112546	0.065615
Std. dev.	0.149978	0.150731	0.141232	0.080698
RMSE	0.149949	0.151559	0.150752	0.080808
C. $n = 500$	β_0	β_1	β_2	ρ
Avg. est.	-0.00512	0.310667	0.827631	0.297883
Avg. bias	-0.00512	0.010667	0.027631	-0.00212
Avg. abs. bias	0.076735	0.07769	0.073923	0.048849
Std. dev.	0.098646	0.099238	0.091461	0.061608
RMSE	0.098729	0.099761	0.0955	0.061613
D. $n = 1,000$	β_0	β_1	β_2	ρ
Avg. est.	0.000227	0.303652	0.813672	0.302332
Avg. bias	0.000227	0.003652	0.013672	0.002332
Avg. abs. bias	0.055414	0.052707	0.048474	0.038313
Std. dev.	0.069681	0.065855	0.060237	0.047887
RMSE	0.069647	0.065923	0.061739	0.047919

Note: The table presents the simulation results for the estimation method that allows for the endogeneity of both the spatial weight matrix and explanatory variables. In the simulation, parameters n denotes the sample size which increases from 100 to 1,000, and we repeat 1,000 Monte Carlo trials for each n, $\beta_0 = 0$, $\beta_1 = 0.3$, $\beta_2 = 0.8$, and $\rho = 0.3$.

Table A.3: Countries in the full sample

Country	Cross-sectional	Panel	Country	Cross-sectional	Panel
Albania	√	√	Ecuador	√	√
Algeria	√	√	Egypt	√	√
Angola	√		El Salvador	√	√
Argentina	√	√	Estonia	√	√
Armenia	√		Ethiopia	√	
Australia	√	√	Fiji	√	√
Austria	√	√	Finland	√	√
Azerbaijan	√		France	√	√
Bahrain	√	√	Gabon	√	√
Bangladesh	√	√	Gambia	√	
Belarus	√		Georgia	√	
Belgium	√	√	Germany	√	√
Benin	√	√	Ghana	√	√
Bhutan	√		Greece	√	√
Bolivia	√	√	Guatemala	√	√
Botswana	√	√	Guinea	√	
Brazil	√	√	Guinea-Bissau	√	√
Bulgaria	√	√	Guyana	√	√
Burkina Faso	√		Haiti	√	√
Burundi	√	√	Honduras	√	√
Cambodia	√		Hungary	√	√
Cameroon	√	√	India	√	√
Canada	√	√	Indonesia	√	√
Cape Verde	√		Iran	√	√
Central African Republic	√	√	Iraq	√	
Chad	√	√	Ireland	√	√
Chile	√	√	Israel	√	√
China	√	√	Italy	√	√
Colombia	√	√	Jamaica	√	√
Congo	√	√	Japan	√	√
Costa Rica	√	√	Jordan	√	√
Croatia	√	√	Kazakhstan	√	
Cyprus	√	√	Kenya	√	√
Czech Republic	√	√	Kuwait	√	√
Denmark	√	√	Kyrgyz Republic	√	
Dominican Republic	√	√	Laos	√	

Note: The table presents the countries in our sample.

Table A.4: Countries in the full sample (continued)

Country	Cross-sectional	Panel	Country	Cross-sectional	Panel
Latvia	√	√	Portugal	√	√
Lebanon	√		Qatar	√	
Lesotho	√		Russia	√	√
Liberia	√		Rwanda	√	√
Libya	√		Saudi Arabia	√	
Lithuania	√	√	Senegal	√	√
Luxembourg	√	√	Sierra Leone	√	√
Macedonia	√		Singapore	√	√
Madagascar	√	√	Slovak Republic	√	√
Malawi	√	√	Slovenia	√	√
Malaysia	√	√	South Africa	√	√
Mali	√	√	South Korea	√	√
Mauritania	√		Spain	√	√
Mauritius	√	√	Sri Lanka	√	√
Mexico	√	√	Suriname	√	
Moldova	√		Swaziland	√	
Mongolia	√		Sweden	√	√
Morocco	√	√	Switzerland	√	√
Mozambique	√		Tajikistan	√	
Myanmar	√		Tanzania	√	√
Namibia	√	√	Thailand	√	√
Nepal	√	√	Togo	√	√
Netherlands	√	√	Trinidad and Tobago	√	√
New Zealand	√	√	Tunisia	√	√
Nicaragua	√	√	Turkey	√	√
Niger	√	√	Uganda	√	√
Nigeria	√	√	Ukraine	√	√
Norway	√	√	United Arab Emirates	√	√
Oman	√	√	United Kingdom	√	√
Pakistan	√	√	United States	√	√
Panama	√	√	Uruguay	√	√
Papua New Guinea	√	√	Venezuela		√
Paraguay	√	√	Vietnam	√	
Peru	√	√	Zambia	√	√
Philippines	√	√	Zimbabwe	√	√
Poland	√	√			

Note: The table presents the countries in our sample.