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The Geography of Conflicts and Regional Trade Agreements†

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In addition to standard trade gains, regional trade agreements (RTAs) can promote peaceful relations by increasing the opportunity cost of conflicts. Country pairs with large trade gains from RTAs and a high probability of conflict should be more likely to sign an RTA. Using data from 1950 to 2000, we show that this complementarity between economic and politics determines the geography of RTAs. We disentangle trade gains from political factors by a theory-driven empirical estimation and find that country pairs with higher frequency of past wars are more likely to sign RTAs, the more so the larger the trade gains. (*JEL* D72, D74, F15, N70)

"…people forget too often about the political objectives of the European constitution. The argument in favor of the single currency should be based on the desire to live together in peace,"
— Jacques Delors (former president of the European Commission 1997).

Regional trade agreements (RTAs) have a bad reputation among a number of economists. Many scholars argue that they constitute a threat to the carefully constructed post-war multilateral trade system. Whereas multilateral trade liberalization has stalled, the number of RTAs has massively expanded during the last two decades, and they are now well over 300. The well-known problem with these bilateral and regional agreements is that, although they create trade, they also generate distortions by excluding countries.[1] Much less attention has been paid (by economists) to the political and strategic motivations for regional integration, even

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[1] The most recent evidence (Baier and Bergstrand 2007, using gravity equations) on trade creation finds a relatively large effect: RTAs are on average responsible for a doubling of trade between two members after ten years. Baier and Bergstrand (2009) use matching techniques and confirm this large effect of RTAs on trade between members. Much less is known about trade diversion and, therefore, the potential economic costs of these preferential agreements. Hence, the economic case for RTAs is still an open debate.

though these motivations may have been key historically.[2] In the case of Europe, political scientists and historians have insisted on the fact that economic integration was viewed as an intermediate objective while its final objective was to prevent the killing and destruction of the two World Wars from ever happening again. Even the recent creation of the euro, often interpreted by economists as a logical step toward more economic integration, has been discussed in these terms (see quote above). Before that, the Cobden-Chevalier Treaty was signed to diffuse tensions between the two countries. Outside Europe, Mercosur was created in 1991 in part to curtail the military power in Argentina and Brazil, then two recent and fragile democracies with potential conflicts over natural resources. In fact, the debate between economists and political scientists often interprets economic and political rationales for RTAs as substitutes. In this paper, we revisit the case for regional integration by explicitly linking the economic and political rationales, and we show empirically that the two complement each other.

An important link between trade policy and conflicts is the so called Liberal Peace argument, which states that bilateral trade flows reduce the probability of a bilateral war, a channel that has been analyzed theoretically and on which some empirical evidence exists.[3] A closely related mechanism is that RTAs, because they create trade, reduce the probability of wars between countries. Combining the estimate of the trade-creating effect of RTAs by Baier and Bergstrand (2007) and the estimate of the trade-induced reduction of war probability by Martin, Mayer, and Thoenig (2008a), we can get a rough estimate of this mechanism. In their preferred regression, Martin, Mayer, and Thoenig (2008a) estimate an elasticity of -0.236 for the impact of bilateral trade openness on war probability for contiguous countries. Baier and Bergstrand (2007) find that an RTA has a large trade creation effect by doubling bilateral trade after ten years. As long as the probability of war between contiguous countries is low (it is estimated at 4 percent in Martin, Mayer, and Thoenig 2008a) this suggests that an RTA, by doubling bilateral trade, would roughly decrease the probability of conflict by 23 percent for a pair of contiguous countries. However, such a back-of-the-envelope estimate should be interpreted with extreme caution. The combination of these estimates above may be biased because of a severe reverse causality issue: two countries may sign an RTA partly because they expect that their conflictuality is going to fade in the near future. In this case, a lower expected probability of conflict causes the signing of RTAs. The lack of historical perspective *following* RTA formation (most RTAs were signed in the 1990s and 2000s) would also make identification difficult in the panel dimension.[4] In this paper, we choose a different route by asking a different question: is the geography of RTAs consistent with a model in which policy makers *believe* that RTAs are pacifying and therefore *believe* in the Liberal Peace argument? This empirical strategy allows us to exploit the period *preceding* RTAs formation for identifying the relevant effects.

[2] Cordell Hull, President F. D. Roosevelt's Secretary of State, was a fervent promoter of free trade with the belief that trade leads to peace.

[3] See for instance Polachek (1980); Oneal and Russett (1999); Martin, Mayer, and Thoenig (2008a); Spolaore and Wacziarg (2009); Hegre, Oneal, and Russett (2010). See also Martin, Mayer, and Thoenig (2008b) for a related argument for civil wars.

[4] Mansfield and Pevehouse (2000) find that country pairs in RTAs are less likely to be in conflict than others. However, their cross-sectional evidence does not allow us to conclude on the direction of causality.

We first use a simple theoretical framework to illustrate the different mechanisms at work in the decision whether to sign an RTA or not. In addition to the expected trade gains, policymakers consider that RTAs provide two types of peace-promoting security gains: by offering a political forum which facilitates settlement of future disputes, and by increasing the opportunity cost of future and potentially trade-disrupting wars. This simple framework allows us to derive several testable implications. First, RTAs are more beneficial to country pairs with a higher probability of war because the expected welfare gain of the political forum channel is larger. Second, expected trade gains and the probability of war have a positive and complementary impact on RTA formation. The complementarity stems from the opportunity cost channel: the larger the trade gains, the larger the opportunity cost of a war, and therefore the more useful an RTA is to secure peace, which is more valuable to countries that have a higher probability of war.

Our empirical analysis estimates a model of RTA formation at the country pair level over the 1950–2000 period to analyze whether the evolving geography of RTAs is consistent with the economic and political factors identified in the theoretical section. From the perspective of the identification strategy, an important concern is that many empirical determinants of wars and of the RTA-related trade gains are confounded: the gravity covariates, such as geographical distance, economic size, contiguity, cultural distance, etc., do affect the propensity to fight and the propensity to trade. This issue explains why the existing empirical literature on RTA formation has difficulty in disentangling the economic factors from the political factors. Here, we propose to rely on a theory-driven estimation procedure to quantify directly the potential trade gains generated by RTAs. To our knowledge we are the first to adopt such a strategy and this is an additional contribution of our paper. We address the various endogeneity issues by controlling for the main codeterminants of political affinity, conflicts and trade; by including various country, country pair, and year-fixed effects; and by instrumenting trade gains. All the results are robust to these different estimation strategies. In particular, we check that the results are not driven by the European integration process although the mechanism is stronger for European country pairs.

Both in the cross-section and in the panel dimension, we find that trade gains and frequency of old wars have a high explanatory power and both increase the occurrence of RTA formation; their interaction term also has a positive impact, and this confirms complementarity between economic and political factors. By contrast, recent war frequency decreases the occurrence of RTA formation, suggesting the presence of windows of opportunity to lock-in RTAs. Periods of interrupted conflict between old enemies may help them to form an RTA in order to settle a more peaceful bilateral relation. Finally, we find that country pairs characterized by multilateral trade openness and a high frequency of old wars are more likely to sign RTAs. We interpret this in the light of one of our main findings in Martin, Mayer, and Thoenig (2008a): Multilateral trade openness, because it reduces bilateral economic dependence and the opportunity cost of a bilateral war, increases the probability that a dispute escalates into a conflict. In other words, countries respond to the weakening of local economic ties (a side effect of multilateral trade liberalization), and its potentially peace-harming consequences, by reinforcing local economic ties through an RTA. From this point of view, we interpret the multiplication of RTAs as a logical political response to trade multilateralism.

In the last section of the paper, we perform a quantitative interpretation of the econometric results. We find that the complementarity between trade gains and the probability of war is sizeable and may even dominate the direct effect of each of these variables. This suggests that trade gains brought by RTAs are instrumentalized and are important as an intermediate objective of RTAs, their final goal being to pacify relations between countries. We also find that in a counterfactual world without any past history of warfare, the geography of RTAs formation would be radically different from the one actually observed. The same is true for a counterfactual world with no multilateral trade openness.

The theoretical economic literature on RTA formation is very large. Nevertheless existing papers focus their analysis on the economic determinants,[5] the role of security gains and military conflicts being largely ignored. From an empirical point of view, several papers study the economic determinants of RTAs (Baier and Bergstrand 2004; Egger and Larch 2008) under the identifying assumption that RTA-related trade gains are closely linked to the standard gravity covariates. As discussed above, this does not allow us to discriminate between the economic and political factors, which is the purpose of our study. Symmetrically, Mansfield and Pevehouse (2000) and Vicard (2012) look at the impact of RTA formation on the occurrence of military conflicts, ignoring the potential role of economic factors.

The next section provides a simple framework and derives several testable implications. Section II presents the data, and discusses the empirical strategy. Section III reports our main empirical results and performs some quantification exercises, while Section IV concludes.

I. A Simple Framework

We now present a cost/benefit analysis of RTA formation. We keep the analysis as simple as possible, our purpose being to derive the minimal setup for grounding our econometric specification (see equation (8)). We leave for future work the building of a fully-fledged theory of the dynamics of RTA formation in presence of trade-disrupting conflicts. Such a theory goes beyond the scope of this paper. Readers who wish to skip this step may go directly to Subsection IC.

A. *Timing and Welfare*

We consider an insecure multi-country world where two countries decide whether to sign a bilateral RTA, which we interpret as a decrease in bilateral trade barriers with respect to the Most Favored Nation (MFN) tariff. We analyze hereafter how this decision is shaped by economic and political forces. For the ease of exposition, we focus, in this section only, on two identical countries.

Two main features describe bilateral relations between countries. First, whether they have signed an RTA or not. The variables of those who have signed an RTA are

[5] This literature has analyzed the motives for building RTAs mainly from a term-of-trade perspective (Bagwell and Staiger 1997; Ornelas 2005) and from a commitment perspective (Limao 2007; Maggi and Rodríguez-Clare 1998).

denoted with a superscript RTA. Those who have not signed have no superscript. The second dimension is whether the two countries are at war or in peace.

The timing of events is as follows: in period 1, countries negotiate on the RTA. We make no particular assumption on the bargaining process but assume that there is a political cost of negotiation C that is borne by each country. In period 2, we assume that a bilateral dispute may arise with probability δ for exogenous reasons (the existence of a common border, natural resources, ethnic minorities,…) and may escalate into a military conflict with an endogenous conditional probability: e in absence of an RTA or e^{RTA} if an RTA is in force. In period 3, trade gains are realized and each country gets an aggregate welfare level, which depends on the existence of an RTA and on the realization of a war at date 2.

In the rest of our analysis, we express all welfare gains or losses as a percentage of a benchmark welfare, U_P, which is realized in the state of peace in absence of an RTA. In this state, both countries trade bilaterally and the MFN tariff level is applied. When war occurs, we assume that bilateral trade is fully disrupted and both countries go back to bilateral economic autarky. This trade disrupting effect of war is empirically well grounded (Blomberg and Hess 2006; Martin, Mayer, and Thoenig 2008a; Glick and Taylor 2010). Hence, welfare under war is given by $(1-W)U_P$ with $0<W<1$, whether an RTA is in force or not. The parameter W captures the direct costs of war (i.e., destructions, death toll, etc.) augmented with the loss associated to bilateral economic autarky (with respect to the MFN situation). When an RTA is in force, additional welfare gains with respect to the MFN situation are generated only if peace is maintained; in that case welfare is given by $(1+T)U_P$. One of the purposes of the empirical analysis is to estimate precisely these trade gains T associated to RTA formation.

The opportunity cost of war corresponds to the welfare differential between war and peace. From the previous discussion we see that in absence of an RTA, this differential is equal to WU_P, while it is equal to $(W+T)U_P$ when an RTA is in force. As a consequence, signing an RTA increases the opportunity cost of a war by T/W percent.

B. *Signing an RTA: Testable Implications*

At date 1, an RTA is signed when, for each country, the expected utility gains induced by the RTA, Γ, are larger than the political cost. Noting V^{RTA} and V the *expected* welfare with and without RTA, the condition for RTA signature is:

$$\Gamma \equiv V^{RTA} - V \geq C, \tag{1}$$

where $V=(1-\delta e)U_P+\delta e(1-W)U_P$ and $V^{RTA}=(1-\delta e^{RTA})(1+T)U_P+\delta e^{RTA}(1-W)\,U_P$. Without loss of generality, we can express the political cost as a percentage of the benchmark welfare: $C=c\times U_P$. Below, we detail some likely determinants of the negotiation cost c. Combining those equations with equation (1), the condition for signing an RTA becomes:

$$\Gamma \equiv \underbrace{(1-\delta e^{RTA})T}_{\text{economic gains}} + \underbrace{\delta(e-e^{RTA})W}_{\text{security gains}} \geq c, \tag{2}$$

where, on the left-hand side, we have decomposed the net expected surplus of RTA formation into pure economic gains and security gains. Economic gains result from the increase in welfare from U_P to $(1+T)U_P$ when the RTA is active. However, the RTA related trade gains T are realized only in periods of peace, which occur with probability $(1-\delta e^{RTA})$. The security gain of an RTA is associated with the potential decrease in the probability of escalation of disputes into war from e to e^{RTA}. This allows us to save on the costs of war W.

We now analyze the differential $(e - e^{RTA})$. As shown by the international relations literature (see Fearon 1995 and Powell 1999 for surveys), escalation to military conflicts can be interpreted as the failure of negotiations in a bargaining game. From this perspective, the probability of escalation depends negatively on the opportunity cost of war and positively on the degree of informational asymmetry between the two countries.[6] The rationale for the first channel is that, as the opportunity cost of war increases, countries have more incentive to make concessions in order to avoid the escalation of a dispute into a military conflict. The rationale for the second channel is that information asymmetries imply that during negotiations, countries do not report their true outside option, in order to extract larger concessions. This may prevent negotiations to succeed and disputes may escalate into war.

The signature of an RTA affects the probability of escalation, e, through these two distinct channels. First, as discussed in the previous section, RTA potentially increases the opportunity cost of war by T/W percent and thus reduces the probability of escalation. Second, regional integration produces a political spillover on conflict resolution by reducing the degree of informational asymmetries. Successful negotiations on economic and trade matters, and the repeated interactions that follow these negotiations, enable policymakers to learn about the other country. This channel has been discussed at length in the political science literature,[7] and many RTAs, such as the European Union (EU), Association of Southeast Asian Nations (ASEAN), or Mercosur, have become venues to discuss political issues and potential disputes. Transposed to our framework, this discussion implies the following assumption on the probability of escalation under an RTA:

$$\frac{e^{RTA} - e}{e} = -\varepsilon_{pol} - \varepsilon_{cost} \times \frac{T}{W} < 0, \tag{3}$$

where $\varepsilon_{pol} > 0$ stands for the political spillover effect, while $\varepsilon_{cost} > 0$ corresponds to the elasticity of escalation e to the cost of war. In the rest of the paper, we refer to $(\varepsilon_{pol}, \varepsilon_{cost})$ as the security gains of RTA formation. This also points to a limitation of our framework: we do not model jointly the decision to sign an RTA and the

[6] For a formal proof, see for example Martin, Mayer, and Thoenig (2008a), where we consider a fairly general bargaining game, such that: war is Pareto dominated by peace, countries have private information on the military and political strength of the other country, and countries can choose any type of negotiation protocol. The negotiation is such that escalation to war is avoided whenever countries agree upon the sharing of the economic surplus under peace.

[7] This argument, under the name of issue linkage, has been developed by political scientists working in the field of international relations, see Keohan and Nye (1977), Haas (1980), and Mansfield and Pevehouse (2000).

escalation to war. In Martin, Mayer, and Thoenig (2008a), we present a model of escalation based on asymmetric information. We leave for further research an analysis where both RTA formation and escalation to war are made endogenous.

Under the reasonable assumption[8] that the RTA-related trade gain T is small with respect to the cost of war W, we can combine (2) and (3) to get an approximation of the RTA signature condition (see the Appendix for the details):

$$\Gamma = T + \varepsilon_{pol}(\delta e \times W) + (\varepsilon_{cost} - 1)(\delta e \times T) \geq c, \tag{4}$$

where Γ corresponds to the utility gains of RTA formation. This equation is our key prediction and will serve as a foundation for the econometric estimation. It contains five main predictions on the determinants of signing RTAs:

- The first term (T) on the left-hand side of this inequality corresponds to the standard trade gains generated by the RTA on which the literature has focused. *Larger trade gains are predicted to increase the probability that the two countries sign an RTA*. The difficulty here is to produce a quantitative estimate of those trade gains for all country pairs. This is what we do in the empirical section.
- The second term corresponds to the positive political spillover of RTAs. *A higher probability of war δe increases the likelihood of signing an RTA*. Because signing an RTA allows to reduce the level of asymmetric information, it reduces the probability of escalation to war by ε_{pol} percent. Note that this political gain of RTAs is large when the potential welfare loss of war W is large.
- The third term interacts trade gains with the probability of war. The sign of this term is ambiguous. It is positive if the elasticity of escalation to the cost of war is suffiiciently large, i.e., if $\varepsilon_{cost} > 1$. Two effects indeed go in opposite directions: on the one hand a high probability of conflict δe reduces the expected gain of an RTA because these gains are lost in times of war. On the other hand, a high probability of conflict also means that the pacifying effect of an RTA is very valuable. If policymakers believe that RTAs are indeed strong elements of pacification, this second effect dominates, and *we expect this interaction term to enter with a positive sign.*[9]
- The c term on the right-hand side is the political cost of negotiation: it is linked to the current state of relations between the two countries.

[8] In the next section, our empirical estimates show that the magnitude of T is approximately 1 percentage point of welfare. This is far below the existing estimates of the average cost of war W that can be found in the empirical literature (see Glick and Taylor 2010).

[9] An alternative interpretation of the interaction term between probability of conflict and trade gains can be provided. The total cost of war can be decomposed in the sum of nontrade-related welfare losses (death toll, destructions, …), Ω, and of trade-related welfare losses (returning to bilateral autarky). This itself could be positively correlated to the trade gains generated by an RTA: αT. With such an assumption, $W = \Omega + \alpha T$, so that equation (4) becomes: $\Gamma \equiv T + \varepsilon_{pol}\delta e \times (\Omega + \alpha T) + (\varepsilon_{cost} - 1)(\delta e \times T) \geq c$. If trade-related welfare losses due to a war are positively correlated to trade gains from an RTA, the interpretation of the interaction term (both from a theoretical and empirical point of view) changes: it represents both the economic opportunity cost of war and the trade gains arising from the "security spillover" of an RTA that improve the bilateral relationship.

- Consider now the effect of multilateral openness. One theoretical and empirical result in Martin, Mayer, and Thoenig (2008a) is that multilateral trade openness reduces the opportunity cost of a bilateral war and, therefore, increases the probability that a dispute escalates into a war. The rationale is that multilateral trade openness provides alternative trade partners and reduces bilateral trade dependence with the countries with which a dispute could escalate. Hence, in the context of our framework, an increase in multilateral openness has two opposite effects. It corresponds to a decrease in the cost of war W and an increase in the probability of escalation e. Assuming that a change in multilateral openness has these effects, we can write the cost of war as $W = (1 - \omega)W_0$ and the probability of escalation as $e = (1 + \varepsilon_{cost}\omega)e_0$ with ω as a measure of multilateral openness, W_0 is the cost of war and e_0 the escalation likelihood when $\omega = 0$. Substituting into equation (4) we obtain

$$T + \varepsilon_{pol}\delta e_0 W_0 + (\varepsilon_{cost} - 1)(\delta e_0 \times T) + (\varepsilon_{cost} - 1)(\varepsilon_{pol} W_0 + \varepsilon_{cost} T)(\delta e_0 \times \omega) \geq c. \tag{5}$$

The coefficient of the interaction term $(\delta e_0 \times \omega)$ is positive when $\varepsilon_{cost} > 1$. Hence, when the elasticity of escalation e to the cost of war is large enough, *multilateral trade openness and the probability of war are expected to have a positive and complementary impact on the probability of RTA formation.* Political motives therefore imply that multilateral trade openness gives an incentive to sign RTAs to country pairs prone to conflict. The intuition is that an RTA is a way to compensate the potentially destabilizing consequence of multilateral trade openness. The incentive to do so increases with the probability of war. This result supports the view that the development of multilateralism during the 1980s and early 1990s could have triggered the wave of regionalism in the late 1990s. This echoes a recent empirical finding by Fugazza and Robert-Nicoud (2011) that in the US case multilateralism has pushed toward regionalism. They indeed find that the extent of post-Uruguay Round RTAs (in terms of included tariff lines) is positively affected by the extent of MFN tariff cuts negotiated by the US during the Urugay Round. While Fugazza and Robert-Nicoud (2011) provide no theory for their intriguing finding, our results suggest that this emulator effect of multilateralism on regionalism could also be driven by security purposes.

C. *Empirical Implementation*

We now present the econometric implementation of our model of RTA formation. To this purpose, we relax the assumption of identical countries. Considering a country pair (i, j) at year t, equation (4) implies that a regional trade agreement is signed when

$$\Gamma_{ijt} > c_{ijt}. \tag{6}$$

In this equation, Γ_{ijt} is the utility gain from signing the agreement, and c_{ijt} corresponds to the negotiation cost. Empirically, c_{ijt} is the unobserved component of the decision process, submitted to stochastic shocks in political affinity for instance, which transforms (6) into a probability of RTA formation. The functional form taken by this probability depends upon the distribution assumed on c_{ijt}. With a Gumbel/Type I extreme value distribution (see Train 2003), we obtain the logit probability to be estimated using maximum likelihood:

$$\mathbb{P}(RTA_{ijt} = 1) = \frac{\exp(\Gamma_{ijt})}{\exp(\Gamma_{ijt}) + 1}, \tag{7}$$

where the dependent variable RTA_{ijt} is a dummy coding for the existence of an RTA between i and j in year t, and Γ_{ijt} follows from equation (4):

$$\Gamma_{ijt} = \alpha + \beta_1 \min(\hat{T}_{ijt}, \hat{T}_{jit}) + \beta_2 WAR_{ij} + \beta_3 \min(\hat{T}_{ijt}, \hat{T}_{jit}) \times WAR_{ij} + \beta \mathbf{Z}_{ijt}. \tag{8}$$

In the previous equation, $(\hat{T}_{ijt}, \hat{T}_{jit})$ correspond to our *empirical estimates* of the RTA-induced trade gains; they are retrieved from the estimation procedure described in Section IIB. We consider the country-pair minimum $\min(\hat{T}_{ijt}, \hat{T}_{jit})$ as a consequence of our assumption that RTA formation must be Pareto-improving in absence of any compensatory transfers within the country pair.[10] In our robustness analysis, we allow for the possibility of transfers by measuring trade gains with the country-pair average $(\hat{T}_{ijt} + \hat{T}_{jit})/2$ rather than the minimum. WAR_{ij} is a proxy for the probability of war (see the discussion in Section IIC on how we measure this variable), $\mathbf{Z}_{ijt}$ is a set of control variables. We include these to filter out potential correlation between our main explanatory variables and the residual term c_{ijt} that captures cross-country pair unobserved heterogeneity in the political costs of negotiation. These controls are discussed below.

In equation (8) we expect β_1 to be positive. The coefficient β_2 tests for the existence of a political spillover of RTA. It is expected to be nonnegative. The interpretation of the sign of β_3, the coefficient of the interaction term, can be misleading in a logit specification due to the nonlinearity of this model (see Ai and Norton 2003). The logit specification also makes the handling of panel data techniques, such as within estimation, more complicated, while the marginal effects tend to be similar to the Linear Probability Model (LPM) in many cases, as shown in Angrist and Pischke (2009, 107). Hence in all specifications of (8), where the interaction term is included, we estimate a linear probability model rather than a logit model. This

[10] In our setup, the two countries, i and j, are assumed to be symmetric for the sake of exposition. Relaxing this assumption and ignoring compensatory transfers, the condition (4) is now country-specific given that the trade gains (T_{ij}, T_{ji}) are potentially asymmetric. An RTA is formed when the minimum of the two country-specific conditions (4) is positive.

standard choice also facilitates the interpretation of the coefficient.[11] In that case, the coefficient β_3 corresponds to a marginal effect and it can be simply interpreted as a test of complementarity versus substitutability between economic and security gains: complementarity ($\beta_3 > 0$) is expected when the opportunity cost channel is at work (i.e., the pacifying effect of RTAs is large so that $\varepsilon_{cost} > 1$).

II. Empirical Analysis

A. *Data*

There are two main parts to the empirical investigations of this paper. In a first step, we estimate the trade gains of RTA formation, which involves essentially running a gravity equation over a sufficiently long time period to be able to identify the trade creation effect of RTA formation in the within country-pair dimension. In a second step, we estimate the econometric model of RTA formation that is exposed in the previous section.

We make use of the gravity dataset constructed for Martin, Mayer, and Thoenig (2008a) and Head, Mayer, and Ries (2010), which is described in greater detail in those two papers. Essentially, any gravity dataset requires source data for a trade flow variable, and a list of gravity controls. The trade flow source is IMF DOTS, with a procedure to extract the most possible information from mirror flow declarations. The list of gravity controls includes the classical bilateral distances, contiguity, colonial linkages, and common (official) language dummies. All those come from the Centre d'Etudes Prospectives et d'Informations Internationales (CEPII) distance database (http://www.cepii.fr/anglaisgraph/bdd/distances.htm). Later in the paper we also use a common legal origin dummy available from Andrei Shleifer at http://post.economics.harvard.edu/faculty/shleifer/Data/qgov_web.xls, and a variable for bilateral genetic distance, available from Spolaore and Wacziarg (2009).

More central in our case are the regional trading agreements: RTA dummy is the dependent variable of our second and main empirical exercice, which explains their formation. RTAs are constructed from three main sources: Table 3 of Baier and Bergstrand (2007) supplemented with information from the WTO website (http://www.wto.org/english/tratop_e/region_e/summary_e.xls) and qualitative information contained in Frankel (1997). The source data for military conflicts is the Correlates of War (COW) project (http://www.correlatesofwar.org/). More precisely, we use the information contained in the Militarized Interstate Disputes database that lists all bilateral interstate conflicts from 1816 to 2001, and quantifies their intensity on a 1 to 5 scale[12] (for a precise description of the source data, see Martin, Mayer, and Thoenig 2008a). We concentrate on the 1870–2001 period because 1870 is essentially the time when most European countries have a stabilized geographical and

[11] However, an area where logit (or probit) is undoubtedly preferable to LPM relates to the predictions one can make when changing one or more variables more than marginally. In that case, probabilities have to be bounded between 0 and 1 by the model in order to yield meaningful predictions. In our quantification exercise, we return to the logit specification.

[12] The scale is the following: 1 = No militarized action, 2 = Threat to use force, 3 = Display of force, 4 = Use of force, and 5 = War, defined as a conflict with at least 1,000 deaths of military personnel.

political structure. As explained below, we will use both an old war and a recent wars variable. The old wars variable is computed as the percentage of years with active military conflicts between two countries during the 1870–1944 period. This creates an immediate problem with countries that did not exist in this period. What is the historical war propensity of the pair Algeria–Nigeria for instance? Due to the absence of detailed information on conflicts for all pairs of ex-colonies and all years prior to independence, we envision several strategies, which range from assuming peace to dropping those observations. Those strategies and results are detailed below in the results section. Recent wars are taken to be the same percentage of military conflicts, but for a moving window of 20 years before the year under consideration. For both variables, we consider only the two most severe types of wars, coded 4 and 5 in the COW database (see Martin, Mayer, and Thoenig 2008a for examples).

In those regressions, there are other bilateral political variables, which serve as controls in the list of RTA determinants. Those include the correlation of roll-call votes recorded for the two countries in the General Assembly of the United Nations (from Gartzke et al. 1999), a dummy for the existence of a military alliance (from COW), and the sum of democracy indices (from Polity IV).

B. *Estimating the Trade Gains of RTA*

The main objective of our empirical analysis is to estimate the econometric model characterized by equations (6), (7), and (8). Nevertheless the first task is to obtain $(\hat{T}_{ijt}, \hat{T}_{jit})$, the estimates for the trade gains of RTA formation between countries (i, j) at date t. The existing literature on RTA formation (Baier and Bergstrand 2004; Egger and Larch 2008) proxies those gains with the standard gravity covariates, such as economic size, geographical distance, remoteness, contiguity, etc., in a reduced-form estimation of RTA formation. Given that our purpose is to understand the relation between economic and political factors, we cannot follow the same route. Indeed it is extremely likely that the gravity covariates affect both economic and political factors. Hence, we rely on a theory-driven empirical strategy to assess the trade gains of RTA formation and to disentangle them from the political factors.

Let us consider the wide class of trade models where aggregate welfare is derived from a CES utility function.[13] We now use subscripts for countries and time.[14] Country i welfare at date t is given by

$$U_{it} = E_{it}/P_{it}, \tag{9}$$

where E_{it} is nominal GDP and P_{it} is the price index. The price index can be written as

$$P_{it} = \left[\sum_k \mu_{kt}\tau_{kit}^{1-\sigma}\right]^{1/(1-\sigma)}, \tag{10}$$

[13] Arkolakis, Costinot, and Rodríguez-Clare (2012) show that a general class of models, where the "import demand system is CES" (among other conditions), exhibit the same welfare effect from a change in trade costs. This class of models includes Krugman (1980) and Anderson and van Wincoop (2003), but also Melitz (2003) and Eaton and Kortum (2002).

[14] One adjustment we must make to our setup when applying it empirically is to account for heterogeneity between different country pairs, and variance in the time dimension.

where σ is the elasticity of substitution between goods, μ_{kt} stands for all factors in the model that makes country k a good exporter,[15] and $\tau_{kit}^{1-\sigma}$ represents bilateral trade freeness, where $\tau_{kit} > 1$ is the iceberg-type price shifter that accounts for all trade barriers. In this context, bilateral trade obeys the following gravity equation governing imports of i from j in year t:

$$m_{jit} = \mu_{jt} E_{it} P_{it}^{\sigma-1} \tau_{jit}^{1-\sigma}. \tag{11}$$

We estimate the welfare gains of an RTA between countries i and j in a partial equilibrium framework. A full-blown general equilibrium analysis should take into account the fact that each envisioned bilateral RTA triggers an endogenous response of supply conditions in all countries of the world. To take a concrete example, in the Krugman (1980) model, $P_{it} = \left[\sum_k n_{kt} p_{kt}^{1-\sigma} \tau_{kit}^{1-\sigma}\right]^{1/(1-\sigma)}$ and p_k are a function of local factor costs, usually assumed to be limited to labor costs, w_k. By changing the worldwide trade costs matrix, each RTA affects not only the two participants, but also third countries: all n_k and w_k are therefore potentially affected. Dekle, Eaton, and Kortum (2007), extended by Ossa (2011) and Eaton, Kortum, and Kramarz (2011), and applied by Caliendo and Parro (2011) to the context of RTA evaluation, have initiated a literature that takes those effects into account. With the exception of Ossa (2011), those papers focus on the case of factor immobility and consider how arbitrary changes in trade costs (RTAs in our case) affect welfare through endogenous changes in both the price index P and wages w for each country in the world. The Dekle, Eaton, and Kortum method shows how to write the model as a system of two sets of N non-linear equations, one for changes in P, the other for changes in w, where N is the number of countries considered. The system is solved numerically, using measured, calibrated, or estimated values of structural parameters of the model, in particular trade elasticities, but also the share of labor in the economy, overall trade balance, and initial trade patterns, depending on the specific version of the model that is used.[16]

Dekle, Eaton, and Kortum (2007) apply this general equilibrium numerical method to rebalancing of trade by 40 countries in 2004; Eaton, Kortum, and Kramarz (2011) look at the effect of a uniform trade cost fall of 10 percent in 1986 on 113 countries; Ossa (2011) considers Nash-type tariff setting for 7 countries in 2004; and Caliendo and Parro (2011) evaluate the impact of North American Free Trade Agreement (NAFTA) from 1994 to 2005. While it is possible to apply Caliendo and Parro (2011) to our case, note that our problem is far more demanding computationally, since we need to contemplate the impact of a potential RTA for *all pairs of countries* for about 50 years. We consider this to be an important endeavor but beyond the scope of the present paper, and prefer to leave it for further research.[17] A second limitation of the partial equilibrium framework we adopt is that we ignore tariff revenues. This generates measurement error in our empirical estimate of the RTA trade gains, which are

[15] In the Dixit-Stiglitz-Krugman model for instance, this term is $n_{kt} p_{kt}^{1-\sigma}$, a positive function of the number of varieties, and negative one of the price charged by firms located in k.

[16] Note that this framework is general enough to be applied to most of the usual trade models, and is very related in this respect to the Arkolakis, Costinot, and Rodríguez-Clare (2012) result of welfare change equivalence between those different models.

[17] Note also that our results are of the same order of magnitude as the general equilibrium estimates, see below.

certainly overestimated. There are two fundamental hurdles that make tariff losses from RTAs difficult to account for in our work. First, we would encounter the general equilibrium issues just described, since tariff revenue changes depend on how each RTA effects supply and demand conditions in all countries of the world. Again, one would need to apply the methodology of Dekle, Eaton, and Kortum (2007) and Caliendo and Parro (2011) to estimate these tariff revenue changes in a general equilibrium framework. Second, even within a partial equilibrium approach, estimating lost tariff revenues would require data on bilateral tariff schedules for the past 50 years. Such data is simply not available, even at a very aggregate level.

Hence, we stick to a framework maintaining analytical solutions that can be brought to the data by considering only the price index, P_{it}, effect of RTA formation, leaving wages in the world unaffected. The level of P_{it} depends on the existence of an RTA through the bilateral trade barriers in equation (10):

$$\tau_{jit} \equiv \exp(-\rho RTA_{jit})\eta_{jit}, \tag{12}$$

where η_{jit} is the residual component of trade costs, while RTA_{ijt} is a dummy variable set equal to one when an RTA is in force between i and j in t. The parameter ρ depends directly on the preferential tariff cut. Importantly, we assume ρ to be the same across all agreements. This is clearly a simplifying assumption. However, we believe it to be a reasonable approximation of long-run effects of RTAs as World Trade Organization (WTO) rules demand the RTA go to a zero tariff bound after ten years (see GATT Art XXIV, Ad Art XXIV and its updates, including the 1994 "Understanding").

We can combine equations (12), (9), and (10) to obtain T_{ijt}, the percentage change in utility of i following an RTA with j:

$$T_{ijt} = \left[\frac{\sum_k \mu_{kt}\eta_{kit}^{1-\sigma}}{\mu_{jt}\exp[(\sigma-1)\rho]\eta_{jit}^{1-\sigma} + \sum_{k\neq j}\mu_{kt}\eta_{kit}^{1-\sigma}}\right]^{1/(1-\sigma)} - 1. \tag{13}$$

We estimate this equation using bilateral trade data over the 1950–2000 period (see Section IIA for the data description). This requires several steps. First, we use our definition of trade costs (12) in the gravity equation (11) to obtain a new version of the gravity equation:

$$\ln m_{jit} = \ln \mu_{jt} + \ln(E_{it}P_{it}^{\sigma-1}) + (\sigma - 1)\rho RTA_{jit} + (1 - \sigma)\ln \eta_{jit}. \tag{14}$$

That can be estimated by a panel specification:

$$\ln m_{jit} = FX_{jt} + FM_{it} + \lambda RTA_{jit} + u_{jit}, \tag{15}$$

where u_{jit} is the error term, FX_{jt} is an exporter $\times$ year fixed effect, and FM_{it} is an importer $\times$ year fixed effect. This specification has the advantage of remaining flexible in terms of the exact underlying trade model, while enabling the extraction of the parameters of interest for the calculation of the utility change in (13). Indeed, comparing (14) and (15), one obtains $\mu_{jt} = \exp(\widehat{FX}_{jt})$, $\exp((\sigma - 1)\rho) = \exp(\hat{\lambda})$, and $\eta_{jit}^{1-\sigma} = \exp(\hat{u}_{jit})$.

Our panel contains bilateral trade flows over the 1950–2000 period. We exploit the within dimension of this dataset, in order to identify $\hat{\lambda}$ from entries and exits into the agreements rather than from a comparison across country pairs. Thus, in (15), we allow u_{jit} to be additively decomposed into a time-invariant and a time-varying element. The regression also includes year dummies. Finally, due to the potential existence of time-varying codeterminants of RTA formation and trade flows in (15), we instrument RTA_{jit} using the contagion index derived by Baldwin and Jaimovich (2009): $\text{contagion}_{jit} = \sum_{k \neq i,j} \textit{export share}_{ikt} RTA_{jkt}$. This index summarizes the threat of trade diversion suffered by country *i* in market *j*, by weighting the count of RTAs signed between *j* and *k* with the share of *k* in *i*'s exports.[18] Our point estimate of $\hat{\lambda}$ is 0.258 (the uninstrumented estimate being 0.311), yielding a predicted increase in bilateral trade of 29 percent from entry into an RTA. For comparison purposes, Baier and Bergstrand (2007), using bilateral fixed effects and year dummies on a panel (for every five years) from 1960–2000, find an estimate of 0.68 (last column of their table 4). Head, Mayer, and Ries (2010) find 0.383 using their tetradic method, which is most comparable with the method used here (none of the cited papers instruments the RTA dummy, however).

Our second step retrieves those point estimates and substitutes them into equation (13). This gives us our empirical estimate of the trade gains of RTA:

$$(16) \quad \hat{T}_{ijt} = \left[\frac{\sum_k \exp(\widehat{FX}_{kt} + \hat{u}_{kit})}{\exp(\hat{\lambda} + \widehat{FX}_{jt} + \hat{u}_{jit}) + \sum_{k \neq j} \exp(\widehat{FX}_{kt} + \hat{u}_{kit})} \right]^{1/(1-\sigma)} - 1,$$

where we use the standard calibration for the elasticity of substitution in the empirical trade literature $\sigma = 5$.[19]

Figures 1 and 2 and Table 1 describe our trade gains variable $\hat{T}_{ijt}$. In Figure 1, we plot the average estimated trade gains of joining an RTA for two types of country pairs: those that do enter a bilateral RTA at some point in our sample, and those that do not. For the second group, we want to make it as comparable as possible to the first one, and therefore, we keep only those country pairs where both members do enter an RTA with a third country but do not sign a bilateral one.[20] The horizontal axis has the number of years before the signature of the bilateral RTA for those who sign it and the number of years until year 2000 for the control group. The difference in trends is clear: the RTA signatories have estimated trade gains that grow as we get closer to the actual signing, whereas nothing visible happens in the control group. This suggests that our measure of trade gains from an RTA can be used as a predictor

[18] Detailed first-stage results available upon request show that the contagion index is an extremely powerful instrument of RTA signatures.

[19] Global Trade Analysis Project (GTAP) version 5, the workhorse model for computable general equilibrium analysis of trade liberalization retains an average estimate of 5.3 (Dimaranan and McDougall 2002). Econometric evidence by Hertel et al. (2007) point to an average elasticity of substitution of 7.0, while Broda and Weinstein (2006) estimate a mean σ of 4.0 for their most recent period and a 3-digit classification (their table IV).

[20] This restriction does not affect radically the shape of the curve. When comparing with the whole set of country pairs, which do not sign a bilateral RTA, the graph looks almost the same.

Figure 1. Utility Gains RTA/No Bilateral RTA

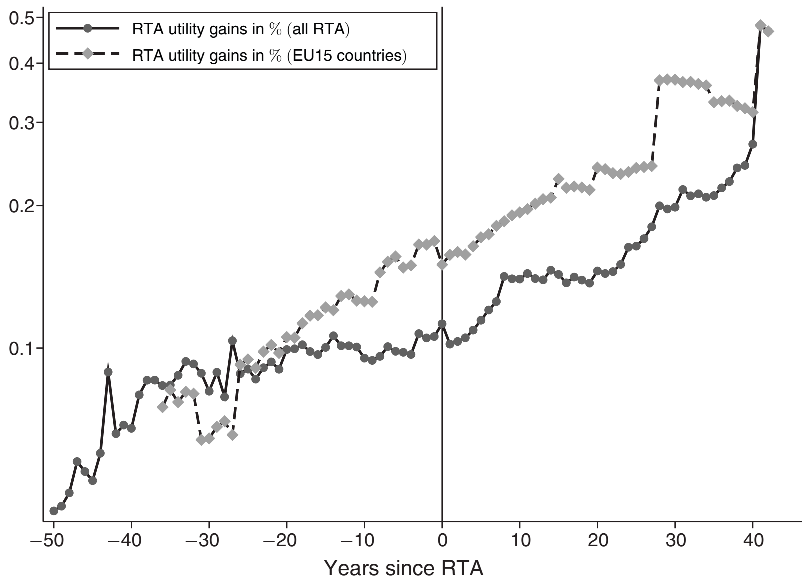

Figure 2. Utility Gains Average RTA/EU15

of the decision to enter a bilateral RTA, both in the cross-section in the years before the signature, and in the within dimension, looking at when countries decide to sign. The variation of $\hat{T}$ across time comes from the changes in the price index decrease that would come with an RTA. As a partner country constitutes an increasing share of your price index, the gain from signing an RTA with this country grows. The source of variance across time and country pairs for T is the variance in bilateral trade barriers η and in exporting country competitiveness μ (see equation 13).

TABLE 1—ESTIMATED TRADE GAINS FOR THE TOP 50 COUNTRY PAIRS IN 1956

Country pair		Trade gains: Minimum T (percent)	Trade gains: Mean T (percent)	Bil. open. $\min \frac{imports}{GDP}$ (percent)	Dist. kms	Ever RTA?
SUN	CHN	1.95	2.919	0.622	5,507	No
USA	CAN	1.786	3.399	0.748	2,079	Yes
NLD	**BEL**	**1.054**	**1.261**	**4.38**	**161**	**Yes**
CZS	SUN	1.031	1.891	0.323	2,388	No
POL	SUN	0.741	1.715	0.231	2,067	No
SYR	LBN	0.667	1.064	2.917	228	No
CAN	GBR	0.637	0.718	1.661	5,850	No
ROM	SUN	0.617	2.294	0.192	2,142	No
FRA	**DEU**	**0.57**	**0.789**	**1.019**	**790**	**Yes**
POL	CZS	0.568	0.701	0.743	387	No
NLD	**DEU**	**0.564**	**0.976**	**1.009**	**379**	**Yes**
GBR	AUS	0.546	1.899	1.128	16,602	No
BEL	**FRA**	**0.546**	**0.754**	**0.559**	**526**	**Yes**
BRA	ARG	0.498	0.555	0.855	2,392	Yes
USA	GBR	0.488	0.713	0.199	6,878	No
USA	BRA	0.469	1.346	0.191	8,089	No
GBR	NZL	0.457	2.165	0.942	18,521	No
USA	VEN	0.444	2.249	0.181	4,204	No
FRA	MAR	0.424	1.986	0.433	1,706	Yes
SUN	FIN	0.385	0.665	0.119	1,635	No
BGR	SUN	0.381	1.84	0.118	2,391	No
BEL	**DEU**	**0.38**	**0.789**	**0.677**	**423**	**Yes**
FRA	IRQ	0.376	0.384	0.383	3,805	No
CZS	CHN	0.369	0.429	0.161	7,790	No
DEU	SWE	0.361	1.017	0.643	929	Yes
USA	JPN	0.352	1.49	0.143	10,286	No
DEU	**ITA**	**0.346**	**0.671**	**0.615**	**1,014**	**Yes**
AUT	ITA	0.338	0.479	0.506	701	Yes
GBR	SWE	0.337	0.702	0.692	1,293	Yes
GBR	IND	0.329	1.161	0.676	7,324	No
GBR	NLD	0.319	0.483	0.657	468	Yes
HUN	SUN	0.319	1.066	0.098	2,334	No
USA	DEU	0.312	0.713	0.127	7,595	No
JPN	PHL	0.301	0.535	0.432	2,957	No
SWE	NOR	0.29	0.676	0.766	503	Yes
USA	CUB	0.289	2.737	0.118	2,581	No
POL	CHN	0.287	0.288	0.125	7,457	No
GBR	DNK	0.285	1.008	0.585	920	Yes
IRN	IND	0.274	0.362	0.235	2,916	No
NLD	**FRA**	**0.274**	**0.276**	**0.284**	**661**	**Yes**
SAU	JPN	0.273	0.315	0.512	8,854	No
ITA	SAU	0.273	0.323	0.408	3,586	No
CHE	DEU	0.273	1.024	0.484	543	Yes
JPN	IND	0.267	0.349	0.372	6,003	No
SWE	DNK	0.266	0.464	0.703	450	Yes
USA	MEX	0.264	2.733	0.107	2,468	Yes
NLD	SWE	0.261	0.402	1.433	1,009	Yes
GBR	FRA	0.261	0.337	0.422	750	Yes
NOR	DNK	0.26	0.263	1.047	560	Yes
CHE	ITA	0.26	0.485	0.388	610	Yes

Note: Lines in boldface indicate pairs that sign the Rome Treaty establishing the European Economic Community a year later.

Figure 2 focuses on the set of countries that do enter an RTA, and distinguishes the European Union members (defined as EU15) from others. We are also able to look at what happens to our measure of trade gains after the RTA signature. One can observe that the trend before signature continues afterwards. This is not surprising: RTA gains come from trade creation, and it is therefore logical that comparing our measure of utility gains before and after the RTA implementation reflects the amount of trade created within the pair. Hence, there is potentially a reverse causality from RTA formation on the trade gains. This points to an important methodological issue that we address in Section IIID.

In Table 1 we report the estimated trade gains in 1956, one year before the Rome Treaty, for the subsample of 50 country-pairs (out of a sample of 8,240) for which the trade gains are the largest. We report the country-pair minimum, $\min(\hat{T}_{ijt}, \hat{T}_{jit})$ and the country pair unweighted average, $(\hat{T}_{ijt} + \hat{T}_{jit})/2$. There may be a large discrepancy between these two figures, especially in asymmetric country-pairs where the smallest country tends to gain much more than the biggest country. In our econometric specifications we focus on the country-pair minimum because it is the theoretically grounded one when welfare transfers and compensation schemes between trade partners are difficult to implement. The interpretation of the table is the following: in 1956, the United States and Canada would have increased their welfare at least by 1.8 percent if they had formed an RTA.[21] Note also that one year before the Rome Treaty, the country pairs composed of the European Economic Community founding countries (in bold) are in the group of large trade winners, but not systematically among the top ones.

The estimated trade gains are small. For instance, in Figure 1, our estimate of the average gain from entering an RTA (at the year of signature) is 0.11 percent. This order of magnitude is not inconsistent with standard results of trade gains estimates based on Computable General Equilibrium (CGE) analysis. A recent example evaluating the impact of the Free Trade Agreement of the Americas by Hertel et al. (2007) finds an estimate average utility gain for potential members of 0.25 percent (their table 5). In the recent general equilibrium literature inspired by Dekle, Eaton, and Kortum (2007), the welfare gains also turn out to be quite modest. Caliendo and Parro (2011) for instance, estimate that the consequence of a counterfactual cancelling of NAFTA would be a drop in welfare of 0.3 percent for the United States, 1.5 percent for Mexico, and 1.4 percent for Canada. Eaton, Kortum, and Kramarz (2011) simulate a 10 percent reduction in all bilateral impediments to trade worldwide in 1986, and obtain gains that are typically within the 0 to 5 percent range (1 percent for the United States, 4 percent for Canada, 3 percent for Germany, and 2 percent for France, for instance).

Finally, our estimate of $\rho(\sigma - 1) = 0.26$ together with the assumption $\sigma = 5$ implies that $\rho = 0.065$. Denoting t the tariff cut following RTA formation, we have $\exp(-\rho) \equiv 1/(1 + t)$. Hence, the estimated average tariff cut is approximatively 6.7 percent for our sample over the 1950–1990 period. This order of magnitude is plausible given, for example, that the difference between MFN and preferential tariff is reported in 2009 by Fugazza and Nicita (2011) to be small, at 2.2 percent.

[21] Regarding this United States-Canada example, the percentage increase in welfare is 1.8 percent for the United Staes and 5 percent for Canada, such as the country-pair average increase is 3.4 percent.

The reason is that in 2009 the MFN tariff had decreased substantially after several decades of multilateral liberalization.[22]

C. *Measuring Conflictuality*

In equation (8), in addition to $\hat{T}$ the trade gains, the second central variable is *WAR*, the probability of war. In order to diminish any issue of reverse causality, we proxy the probability of war at date t, WAR_{ijt}, with the country-pair frequency of bilateral wars that occurred between 1870 and 1945. We call it frequency of *old wars*. This proxy being time invariant, we suppress the time index, which gives the variable WAR_{ij} in the econometric equation (8). The frequency of old wars, WAR_{ij}, therefore proxies for the bilateral probability of war (δe in equation (4)). We also control for the frequency of *recent wars,* which we view as both increasing the political cost of RTA negotiation c and being correlated with WAR_{ij}. This should alleviate part of the correlation between old wars and the residual. Recent wars are very likely to be related also to δe, since empirical evidence suggests that bilateral war probability is stable over time (Collier, Hoeffler, and Söderbom 2004). In order to identify δe and c, we need recent wars to contain additional information about c, while controlling for δe in the regression through the inclusion of old wars. The identifying assumption is therefore that recent wars raise the political cost of subsequent bilateral negotiations, but this effect on the cost then decreases over time. One way to think about this is that feelings of revenge and grievance that follow a war are most vivid just after a war, and then "depreciate" over time (see Mocan 2008 in a different context). For recent wars, we use the country-pair frequency of bilateral wars that occurred during the last 20 years. In our robustness analysis we test definitions of old and recent wars with alternative time spans.

D. *Endogeneity Issues*

The estimates of our main coefficients of interest, β_1, β_2, β_3, are potentially contaminated by several sources of endogeneity, which we now discuss.

Measurement Error: Relying on the old history of conflicts helps to reduce reverse causality issues (from RTA negotiation to war) but introduces noise in the measurement of currently relevant war probability. Some causes of disputes in the late nineteenth century (e.g., the building of colonial empires) may have lost their explanatory power. Simultaneously, new causes have emerged in the late twentieth century (oil production; access to water supply; religious tensions). Those time-varying determinants imply measurement error in the current probability of war. This should go against our results by inducing a bias toward zero in the estimated coefficients of interest.

As discussed in Section IIB, our estimate of the trade gains, $\hat{T}_{ijt}$, relies on a partial equilibrium analysis and is therefore a noisy measure of T_{ijt}, the true trade gains. This

[22] We thank an anonymous referee for suggesting to us this computation.

could bias β_1 and β_3, the two coefficients involving $\hat{T}_{ijt}$, toward zero, underestimating the magnitude of the effects we are trying to identify. However, we see no particular reason why the measurement error $(\hat{T}_{ijt} - T_{ijt})$, which is part of the error term, should be correlated with our other variable of interest, WAR_{ij}, which should leave the estimate of β_2 unaffected by this measurement error problem.

Reverse Causality: Figure 2 highlights the possible reverse causality link from RTA to trade gains *following* RTA formation. In order to eliminate this issue that can overestimate the coefficient β_1, we need to compare $\hat{T}_{ijt}$ across country pairs or time *before* the agreement actually takes place. Similarly, this reverse causality issue may bias downward β_2 because RTA formation is likely to reduce the probability of future conflicts. In the cross-section dimension, we thus estimate equation (8) in year $t = 2000$ for dyads where an RTA does not exist in 2000. For dyads where the two countries are members of an RTA in 2000, their RHS variables are set to their values one year before the RTA formation. For example, in the case of United States-Canada, this means that all the RHS variables take their 1988 values. This methodology generalizes the approach by Baier and Bergstrand (2004) and allows us to control for reverse causation. Correspondingly, in the panel estimates of (8), we focus on "RTA onset," that is we analyze, for each dyad, years up to the signature of the RTA, dropping observations after the signature. This is very similar to the method used by researchers studying the determinants of conflicts (Fearon 2005).

Omitted Variables: In equation (8), the coefficients of economic gain and of its interaction term with the probability of war, β_1 and β_3, could be contaminated by omitted co-determinants of trade gains, $\hat{T}_{ijt}$, and of unobserved political costs of RTA formation, c_{ijt} (i.e., the residual). This may arise because the structural relationship (16) defining $\hat{T}_{ijt}$ depends on $\hat{u}_{jit}$, the estimate of (logged) bilateral trade freeness retrieved from the auxiliary gravity equation (14). Indeed, several determinants of bilateral trade freeness (or conversely trade barriers) might also affect the bilateral political affinity, and consequently the political costs of RTA formation (e.g., commonality of language and culture, economic embargo, etc.). A striking illustration is provided in Michaels and Zhi (2010), who show that the deterioration of political relations between the US and France over the 2002–2006 period resulted in a significant increase in their bilateral trade barriers following changes in attitudes in the United States toward France.

To address this concern, we first add to the set of control variables $\mathbf{Z}_{ijt}$ a series of codeterminants of bilateral trade barriers and political relations. This encompasses the standard time invariant gravity controls (distance, contiguity, common language, etc.) and various time-varying proxies of bilateral political affinity, such as a dummy variable coding for the existence of a military alliance, a measure of bilateral correlation in UN votes from Gartzke et al. (1999), and lastly the country-pair sum of democracy indices from the Polity IV database. Indeed, the democratic peace hypothesis, which has been studied by both political scientists and economists (see Levy and Razin 2004 for a recent explanation of the hypothesis) states that democratic countries are less prone to violence. But democratic countries are also more open to trade. In the panel specifications, we can be more general in those controls, by including a country-pair fixed effect to purge from remaining time-invariant unobserved heterogeneity.

In spite of all these controls, we cannot rule out the possibility that the coefficient of trade gains, β_1, is still contaminated by unobserved *time-varying* codeterminants of bilateral trade freeness, $\hat{u}_{jit}$, and political affinity, c_{ijt}. To solve this last problem, we directly include $\hat{u}_{jit}$ as a control variable. This strategy allows us to identify β_1 by exploiting the variations in trade gains $\hat{T}_{ijt}$ net of $\hat{u}_{jit}$. This solves the omitted variable problem because those variations are not driven by bilateral shocks and so cannot be correlated with the (residual and unobserved) political costs of negotiations ϵ_{jit}. Indeed, a look at the structural relationship (16) makes it clear that those variations are driven by changes in the exporter fixed effects $\widehat{FX}_{kt}$. This strategy is in fact akin to a control function approach where the trade gains $\hat{T}_{ijt}$ are instrumented with a remoteness index based on the exporter fixed effects $\widehat{FX}_{kt}$ (see Imbens and Wooldridge 2007).

Regarding β_2 and β_3, the coefficients of the probability of war and of its interaction term with trade gains in the econometric specification (8), the omitted variable problem is potentially severe. Any time-invariant determinant of the unobserved political costs of RTA formation c_{ijt}, is also likely to affect the underlying probability of war, WAR_{ij}. For example, disputes linked to common borders, natural resources, migration waves, etc., are likely to increase the underlying probability of war and make negotiation on RTA formation politically more costly. This suggests that the omitted variable problem should induce a downward bias which goes against our hypothesis. Note that the various gravity and political affinity controls included in $\mathbf{Z}_{ijt}$ are likely to absorb most of the cross-sectional variations in bilateral disputes. We also include, as a control variable, a measure of bilateral genetic distance. Spolaore and Wacziarg (2009) show that genetic relatedness has a positive effect on bilateral conflict propensities in the cross-section. This is because more closely related populations, on average, tend to interact more and develop more disputes over sets of common issues. Hence, we expect genetic distance to reduce the probability of war and to increase the probability of RTA formation. Finally, in our panel estimates, we include country-pair fixed effects. This makes impossible the identification of β_2, the coefficient of the time-invariant variable WAR_{ij}. Nevertheless, we can still estimate β_3, which is now immune to the omitted variable bias. There is no particular reason for the determinants of political costs c_{ijt} to have a larger effect on RTA formation in dyads where trade gains are larger.

III. Results

A. *Econometric Estimates*

We start in Table 2 with a cross-sectional analysis of RTA determinants. By cross-sectional we mean that we take the world in the year 2000, and attempt to explain which of the country pairs are in an RTA. Some determinants will be time invariant (e.g., distance), some will have a time dimension. For the latter set of variables, we consider the variable for the year immediately preceding the signature of the RTA. For instance, trade gains are taken in 1956 (the year before the Rome Treaty) for the Franco-German case, and in 1993 (the year before NAFTA) for the United States-Mexico

TABLE 2—RTA DETERMINANTS, BENCHMARK REGRESSIONS

Model Dependent variable	RTA (1)	RTA (2)	RTA (3)	RTA (4)	RTA (5)
Period	2000	2000	2000	2000	2000
Trade gains ($\hat{T}_{ijt}$)	0.589*** (0.038)		0.553*** (0.038)	0.415*** (0.019)	0.330*** (0.021)
War frequency pre-1945 (WAR_{ij})		11.716*** (1.104)	7.840*** (1.337)	8.328*** (1.271)	9.257*** (1.273)
Dyad did not exist pre-1945				−0.783*** (0.107)	−0.902*** (0.109)
ln bil. trade freeness					0.217*** (0.025)
Method	Logit	Logit	Logit	Logit	Logit
Sample		Pre-1945 pairs		Whole	Whole
Observations	1,694	2,042	1,694	9,836	9,836
R^2	0.241	0.065	0.263	0.224	0.240

***Significant at the 1 percent level.
**Significant at the 5 percent level.
*Significant at the 10 percent level.

one. Since this variable is calculated as a percentage of utility, it is relevant at the moment of the decision, and can be compared across observations.

Our first column is a logit with only the log of the estimated trade gains as an explanatory variable for RTA formation.[23] Its coefficient is positive and significant. With an R^2 around 0.24 in the context of a logit specification, the explanatory power of this variable is very large. The fact that this variable alone is sufficient to explain more than a quarter of the observed variance in RTA formation provides encouraging empirical support to our theory-driven estimate of trade gains. In the second column, we estimate a similar specification with our second main explanatory variable only, namely the frequency of old wars. Again the coefficient is positive, very significant, with a high explanatory power (i.e., R^2 close to 0.06). In column 3, both variables are included; their respective coefficients and statistical significance remain stable with respect to the estimates of the first two columns. In columns 1–3, the old war variable WAR_{ij} is restricted to the small number of dyads which existed before 1945. In particular, all country pairs that involve a former colony (India-Japan, Germany-Ivory Coast for instance) are dropped from this regression. In column 4, we adopt the following alternative strategy: we set WAR_{ij}, the old war variable, to 0 for country pairs that did not exist before 1945; we also include a dummy variable coding for those pairs. As can be seen from the comparison of columns 3 and 4, the two variables of interest have very close coefficients with this procedure and the fit is very comparable, which makes us confident that it does not alter our results while substantially augmenting the number of observations.[24] We maintain this procedure throughout.

[23] We take the log because of the left-skewness of the distribution of estimated trade gains.
[24] It can be noted that those nonexisting dyads, mostly combinations of colonies at the end of WWII, have been less involved in the RTA movement, as revealed by the negative coefficient of the dummy variable.

Column 5 introduces $\hat{u}_{jit}$, the estimate of bilateral trade freeness obtained from the gravity equation (14). As stated above, this is intended to circumvent any contamination of the coefficient on trade gains, by unobserved codeterminants of bilateral trade freeness and political affinity. As expected, this variable enters positively and results in a decrease of the effect of trade gains as it purges from contemporaneous bilateral affinity, which causes both the probability of signing an RTA and the trade gains to be high.

One of our main variables of interest is the interaction term between old wars and RTA trade gains. Interaction terms have a nonstraightforward interpretation in discrete choice models like the logit, because of their nonlinear nature (Ai and Norton 2003). As explained in details above, we therefore resort to a linear probability model (LPM), which has the additional advantage of handling fixed effects more easily in our panel estimates. Results are in Table 3. Column 1 of Table 3 is simply the LPM version of the logit specification (column 5 of Table 2). While this different estimation method naturally yields different coefficients, the signs and significance levels are preserved. Column 2 introduces the interaction term of trade gains with old wars. This interaction term enters positively and is significant at the 1 percent level. This supports our hypothesis that trade gains and security gains are complements. Dyads with large estimated trade gains are more likely to enter an RTA, and this effect rises with the historical intensity of wars between the partners.

Column 3 tests our equation (5), namely that multilateral trade openness and the probability of war have a positive and complementary impact on the RTA decision. As expected, the coefficient of the interaction term between multilateral openness and old war is positive and highly significant. In column 3, we also include a number of bilateral controls. The two most important gravity variables are geographical distance and contiguity. We also add a list of controls for political affinity (UN vote correlation, the sum of Polity IV reported democracy indices, a dummy for the existence of a military alliance and an index of genetic distance). All of those variables add to the likelihood of belonging to the same agreement as expected. We also include the frequency of recent wars, which, according to our discussion in Section IIC, is expected to enter negatively through their effect on the political cost of negotiations. The coefficient is negative and significant at the 1 percent threshold. The opposite signs of the old wars and recent wars coefficients suggest that a "window of opportunity" mechanism is at work. Having had a history of conflicts in the past makes a country pair more likely to sign an RTA at the condition that their recent history is not too conflicting: Any exogenous event that prevents two ancient enemies to fight for some period improves the chances that they sign an RTA, with the consequence of reducing further the chances of conflict escalation. We quantify the size of those effects later in the paper.

Column 3 establishes our main results with a substantial set of controls, and we consider it as our benchmark specification. In spite of the inclusion of all of these control variables and the resulting reduction by one-third of the sample size, all five coefficients of interest in column 3 keep the expected sign and remain statistically significant at the 1 percent threshold. According to our discussion in Section IC, the fact that the coefficients of the two interaction terms are both positive confirms the internal consistency of our setup.

Table 3—RTA Determinants, Benchmark Regressions, Continued

Model Dependent variable	RTA (1)	RTA (2)	RTA (3)	RTA (4)	RTA (5)
Period	2000	2000	2000	1950–2000	1950–2000
Trade gains ($\hat{T}_{ijt}$)	0.016*** (0.001)	0.014*** (0.001)	0.009*** (0.001)	0.007*** (0.001)	0.002*** (0.000)
War frequency pre-1945 (WAR_{ij})	1.912*** (0.116)	5.169*** (0.341)	7.963*** (0.757)		
Dyad did not exist pre-1945	−0.053*** (0.007)	−0.063*** (0.007)	−0.045*** (0.008)		
ln bil. trade freeness	0.010*** (0.001)	0.010*** (0.001)	−0.011*** (0.002)	−0.003*** (0.001)	−0.002*** (0.000)
Trade gains × wars pre-1945		0.354*** (0.035)	0.460*** (0.042)	0.163*** (0.026)	0.062*** (0.011)
War frequency $[t-20; t-1]$			−0.441*** (0.076)	−0.067*** (0.019)	−0.003 (0.008)
ln distance			−0.127*** (0.004)		
Contiguity			0.095*** (0.018)		
UN vote correlation			0.076*** (0.011)	0.054*** (0.005)	0.003 (0.002)
Sum of democracy indexes			0.053*** (0.006)	0.024*** (0.003)	0.011*** (0.001)
Military alliance			0.127*** (0.010)	0.110*** (0.007)	0.027*** (0.003)
Genetic distance			0.009** (0.003)		
Multi. openness			−0.021*** (0.004)	0.004** (0.002)	−0.001 (0.001)
Multi. openness × wars pre-1945			1.567*** (0.292)	0.629*** (0.053)	0.195*** (0.024)
Method	LPM	LPM	LPM	Cty pair FE	Cty pair FE
Sample	Whole	Whole	Whole	Whole	Whole
Observations	9,836	9,836	6,152	36,701	35,737
R^2	0.138	0.147	0.359	0.081	0.017

***Significant at the 1 percent level.
**Significant at the 5 percent level.
*Significant at the 10 percent level.

The two remaining columns extend the sample to the panel dimension. Both specifications include country-pair fixed effects. The coefficient on old wars cannot be estimated any more, but its interaction with trade gains can. For each dyad, we average data over nonoverlapping time windows of five years, a method comparable to Egger and Larch (2008) and Martin, Mayer, and Thoenig (2008a) in related work. Column 4 considers the full sample. In column 5, we drop observations following the signature of RTA for those who do become members. This RTA onset specification is very demanding and, in spite of the five-year averaging procedure, it is highly sensitive to measurement errors in the time-series dimension. With respect to the benchmark cross-sectional estimates in column 3, all the coefficients of interest keep their expected sign and are statistically significant, with the exception of the coefficient

Table 4—RTA Determinants, Robustness

Model Dependent variable	RTA (1)	RTA (2)	RTA (3)	RTA (4)	RTA (5)	RTA (6)	RTA (7)
Trade gains ($\hat{T}_{ijt}$)	0.296***	0.007***	0.003**	0.005***	0.005***	0.004***	0.005***
	(0.042)	(0.001)	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)
War frequency pre-1945 (WAR_{ij})	44.866***	8.209***	4.637***	6.175***	6.075***	3.823***	6.046***
	(15.989)	(0.754)	(0.900)	(0.670)	(0.662)	(0.676)	(0.662)
Trade gains × wars pre-1945	1.582	0.463***	0.302***	0.333***	0.324***	0.193***	0.325***
	(1.003)	(0.041)	(0.071)	(0.037)	(0.037)	(0.037)	(0.037)
War frequency $[t-20; t-1]$	−7.423***	−0.464***	−0.500***	−0.188***	−0.173**	−0.154**	−0.321***
	(2.123)	(0.076)	(0.081)	(0.069)	(0.068)	(0.066)	(0.106)
Multi. openness × wars pre-1945	17.364***	1.684***	0.777 **	1.446***	1.396***	0.865***	1.375***
	(5.980)	(0.291)	(0.325)	(0.257)	(0.254)	(0.253)	(0.254)
Multi. openness	−1.995***	−0.020***	-0.027***	−0.222***	−0.216***	−0.218***	−0.217***
	(0.233)	(0.004)	(0.005)	(0.012)	(0.012)	(0.011)	(0.012)
Number of landlocked in dyad		−0.000	−0.004	−0.850***	−0.855***	0.044	−0.852***
		(0.005)	(0.006)	(0.163)	(0.161)	(0.156)	(0.161)
Common language		−0.019**	-0.012	−0.020***	-0.014*	−0.012	−0.015*
		(0.008)	(0.009)	(0.008)	(0.008)	(0.007)	(0.008)
Colonial link		−0.031	−0.029	−0.075***	−0.066***	−0.052***	−0.066***
		(0.019)	(0.020)	(0.017)	(0.017)	(0.017)	(0.017)
Common legal origin		−0.002	−0.010	−0.003	-0.017***	−0.011**	−0.017***
		(0.006)	(0.007)	(0.005)	(0.005)	(0.005)	(0.005)
Remoteness		0.083***	0.097***	−0.126***	−0.153***	−0.080***	−0.152***
		(0.011)	(0.013)	(0.025)	(0.025)	(0.024)	(0.025)
Same region					0.114***	0.050***	0.116***
					(0.010)	(0.010)	(0.010)
War frequency $[t-40; t-20]$							0.16*
							(0.092)
Method	Logit	LPM	LPM	Cty FE	Cty FE	Cty FE	Cty FE
Sample	Whole	Whole	Whole	Whole	Whole	No EU15	Whole
Trade gains	Min	Min	Average	Min	Min	Min	Min
Observations	6,152	6,152	5,274	6,152	6,152	6,071	6,152
R^2	0.576	0.366	0.350	0.572	0.582	0.518	0.582

***Significant at the 1 percent level.
**Significant at the 5 percent level.
*Significant at the 10 percent level.

on new wars in the RTA onset specification. An important change is also the size of the coefficient on trade gains, when going from RTA (in column 3) to RTA onset (in column 4) as a dependent variable. This was to be expected from our analysis of Figure 2 and from our discussion of the reverse causality issue: RTAs boost trade volumes, which reinforces the RTA-related trade gains after their implementation.

Table 4 pushes further the robustness investigation. Those regressions take column 3 of Table 3 as a benchmark specification (with gravity controls unreported). In the first column, we re-estimate this benchmark specification using logit instead of LPM. All signs of the relevant variables remain unchanged. The global explanatory power is very high, and the level of significance of the interaction term between old wars and trade gains is now slightly above 10 percent (11.5 percent exactly). This logit estimate is the one that we use in the quantification section.

In the second column, we return to LPM and extend the set of gravity controls to include common language or legal system, colonial linkages, landlockness, and remoteness of the country pair. All our variables of interest keep the same sign. Column 3 changes the definition of bilateral trade gains to be the average of the two countries RTA-related trade gains rather than the minimum. Given that the minimum is always smaller than the average, this translates mechanically into a decrease in the coefficient of trade gains.

Column 4 adds a set of dummy variables coding for each country, a feature which can be properly identified in our cross-sectional sample of (nondirectional) country pairs. These dummy variables control for all time-invariant unobserved characteristics of a country that might make it more likely to fight wars in the past and to sign RTAs now. The global fit naturally increases substantially while leaving our results of interest remarkably similar. Column 5 adds a dummy to control for the fact that the two countries belong to the same geographical region of the world (following the World Bank definition of regions). This increases the probability of RTA significantly, while again leaving our results on trade gains and conflictuality unaffected.

Column 6 removes intra-EU observations by excluding all country pairs where both countries belong to the European Union at 15. This is intended to check that our results are not entirely driven by European countries, which are characterized both by a rich history of warfares and by the creation of the worldwide deepest trade agreement. In this specification, all variables related to wars have slightly—and unsurprisingly—smaller coefficients, but they remain very significant.

Column 7 extends our definition of old wars by including a variable that accounts for war frequency 20–40 years before RTA signature. This results in a smoother representation of the history of wars with very recent ones, those that are more than one generation old, and the very old wars (before 1945). The pattern of coefficients is that recent wars tend to reduce the RTA probability, less recent ones tend to slightly promote them, while old wars have a much stronger positive effect. This finding matches well with our identification strategy: The political and subjective costs imposed by recent wars during RTA negotiation are gradually overturned by the positive strategic effect of war history. In unreported specifications, we test other cutoffs for recent versus old wars; the results were qualitatively unaffected. The same is true for a specification where we account for the full history of past wars (and not only wars before 1945) without controlling separately for recent wars.

B. *Quantification and Counterfactual Experiments*

Up to this point, we have mostly analyzed the signs and statistical significance of coefficients. We now want to quantify the magnitude of the effects we have identified. In order to calculate counterfactuals we need to resort to a logit econometric model where the RTA probability cannot go outside the 0–1 range. Moreover, the presence of interaction terms, which are the core of our analysis, are not straightforward in this context.

In all that follows we adopt the following strategy. We start by running a benchmark regression using logit (column 1 of Table 4) to estimate the coefficients of interest, which gives us the benchmark probability of signing an RTA for each country pair in the sample. We then select a group of observations, and we run a counterfactual by attributing to them other values for one or more explanatory variables. For instance, we take the country pairs in the lowest decile of the frequency of the old war variable and we give them an artificial history of wars. Using the logit formula with the benchmark estimated coefficients, we recalculate their RTA probability and compare it with the benchmark probability to evaluate the magnitude of the effect of the altered variable. This procedure ensures that the probability remains in the admissible range, while doing a "what if" experiment. What if low conflict dyads had had an intense past history of warfare, keeping everything else constant?

Complementarity is a First-Order Effect.—We first quantify our complementarity result between old wars and trade gains in the formation of RTAs. The coefficient of the interaction term between trade gains and old wars is positive both in our benchmark LPM specification (column 7, Table 2) and in our benchmark logit specification (column 1, Table 4). However, Ai and Norton (2003) show that interaction terms have a sign that can be deceptive in a logit framework and that cannot be interpreted readily. In the Appendix, we investigate this question more fully by calculating the marginal effect of this interaction term for the whole range of benchmark probabilities. This confirms that the interaction term is indeed positive for nearly all of the sample.

We now turn to the quantification of the interaction term. To this purpose, we choose pairs of countries that are located inside the middle decile of those two variables, that is, around the median level of old wars and trade gains. We then calculate the ratio of counterfactual to benchmark probabilities of RTA formation following the procedure just described, and spanning over the tenth to the ninety-fifth percentiles of each variable. Results are shown in Figure 3.

In panel A, it is clear that trade gains increase the probability of signing an RTA, and that the effect is amplified with old wars. Panel B allows us to better illustrate the effect. The x-axis reports trade gains while the y-axis reports the ratio of counterfactual to benchmark probabilities. Each curve corresponds to different levels of old wars. For a dyad that moves from the median to the top 20 percent of trade gains, the RTA probability is multiplied by two (1.96) if the dyad is in the middle range of old wars, while the multiplicative factor is almost 3 if the same dyad is in the top 10 percent of war history.[25] We see that the interaction term has a first-order importance. This confirms our intuition that trade gains are important mostly because they allow for an increase in security gains from RTA formation.

Windows of Opportunity.—Our second simulation uses the same method described at the start of this section to quantify the existence of windows of opportunity during which interrupted conflict between old enemies may help sign an RTA and "lock in" a more peaceful bilateral relation. The left panel of Figure 4 is very similar to the one in Figure 3. We take the whole set of dyads with no history of recent or old wars, and gradually move them into the war space, looking at the changes in RTA probability. As expected from the point estimates in Tables 2 and 3, recent wars reduce the probability of RTA formation, while old ones increase it. The magnitude of the effects is substantial. Panel B uncovers an interesting trade off that leaves the

[25] The benchmark probabilities of signing an RTA in this precise sample have an average value of 7.7 percent. The median is much lower at 0.75 percent, which shows that most country pairs in the world have a very low RTA probability, while a few of them have quite a high one (10 percent of the sample has a benchmark probability higher than 20 percent).

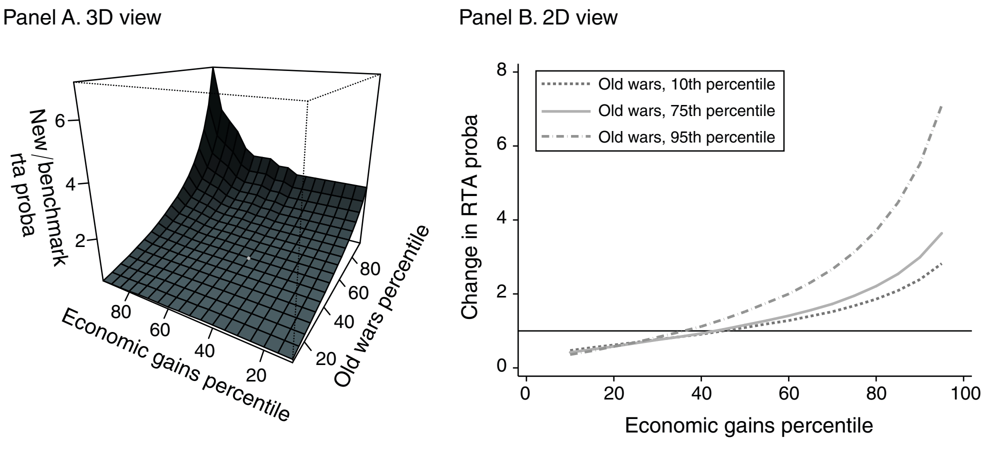

Figure 3. Complementarity between Trade Gains and Security Gains

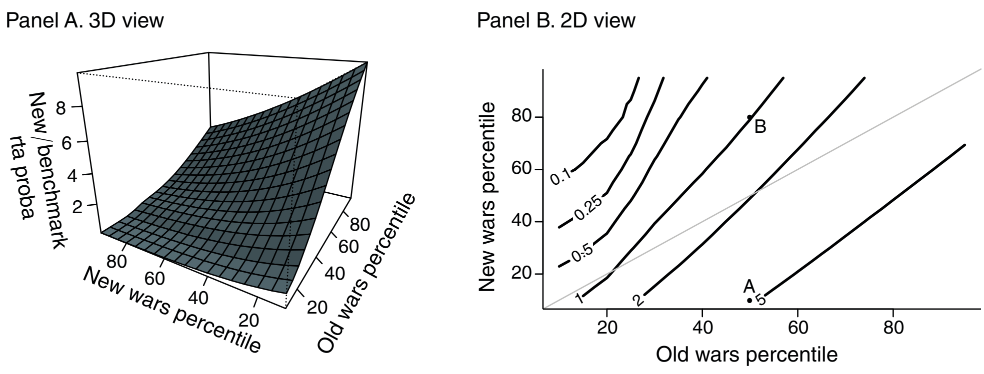

Figure 4. Old Wars and New Wars

change in RTA probability unchanged. Panel B is a contour plot, where each curve represents a probability ratio from panel A. Old wars are on the *x*-axis, recent wars are on the *y*-axis. Assume a country goes from no old wars to the median level. This multiplies its benchmark RTA probability by almost 5 (point A in the figure) if there has been very few recent wars, while it leaves the probability unchanged if the level of recent wars moves to the top 20 percent (point B in the figure). This shows that a change in old wars has in general a larger effect than an equivalent change in recent ones (as revealed by the 45 degree line). In other words, if a country pair's recent history of warfare perfectly reflects its long run history, then the net, overall effect of war is to increase the probability of RTA formation. By contrast, suppose now that we assign the top 5 percent level of old wars to a country pair with no old wars. This multiplies by 10 its RTA probability if recent wars are very rare, but only by 3.5 if the country is also in the top 5 percent of recent conflicts.

We analyzed pairs that did not experience any conflict in the real world. In Figure 5 we take the opposite focus, and look at the effect of recent wars on country pairs that experienced a large set of conflicts in the recent years. We consider four different dyads: India-Pakistan, Greece-Turkey, Egypt-Israel, and Iran-Turkey. Out of those,

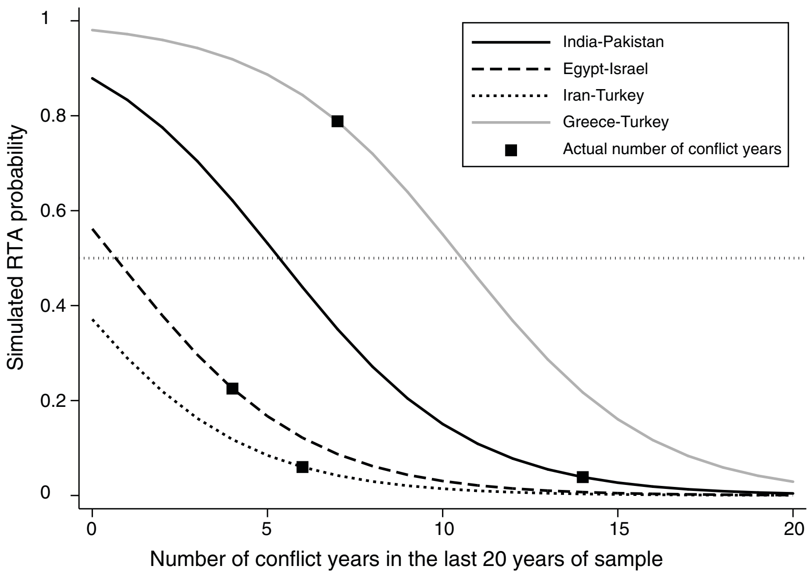

Figure 5. Window of Opportunity for Four Emblematic Country Pairs

Greece and Turkey are the only ones in an RTA (through the customs union signed between the EU and Turkey in 1996).[26] For those four pairs, our variable measuring the proportion of recent conflicts (over the last 20 sample years) spans from 20 to 70 percent (4 to 14 years), with associated benchmark probability ranging from 4 to 80 percent as represented by the black squares on the graph. We then change the number of recent conflict years and calculate the new RTA probability. India-Pakistan is perhaps the most impressive example. After five years of peace, the RTA probability is multiplied by 5 at 20 percent, after 10 years it jumps to 62 percent. Our results also reveal that 4 years of peace between Egypt and Israel brings their RTA probability from 23 percent to 57 percent. The effect of recent wars is quite abrupt for pairs that fundamentally have a large RTA signature probability (those with large potential trade gains, high proximity, …). It thus suggests that the window of opportunity argument may be well grounded. For those pairs, even a short interruption of outbreaks in conflicts can increase RTA probability to a large extent and start a virtuous pacifying process. For Greece-Turkey, we observe the same overall shape of the impact of recent conflictuality, and note that in 1996, the conflictuality between the two countries seemed to have fallen to a level that made RTA possible.

A World without Wars.—Let us consider now another counterfactual experiment. Instead of taking the peaceful dyads and making them fight, we make every country pair peaceful. The frequencies of old wars, recent wars, and all their interaction terms are set to zero, and the resulting, counterfactual probabilities of RTA formation are estimated. Results are reported in Figure 6, where the benchmark probability

[26] The recent war frequency variable is therefore calculated for 1976–1996 for Greece-Turkey, and for 1980–2000 for the three other pairs.

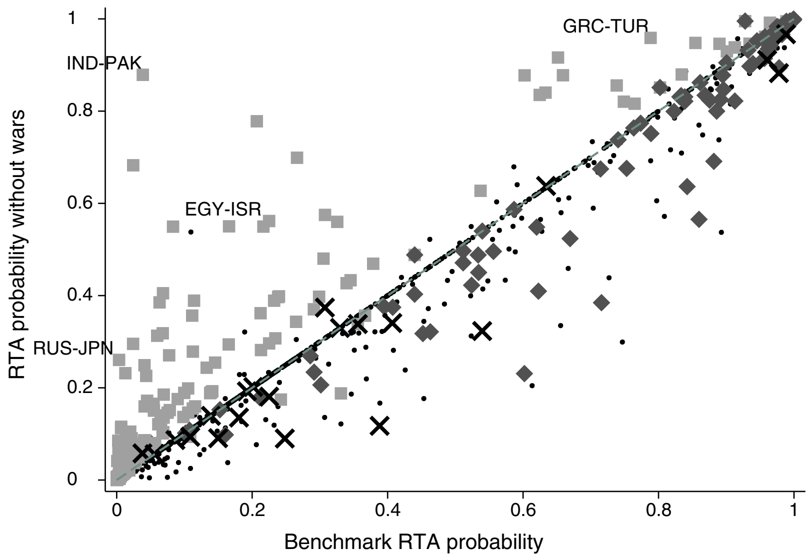

Figure 6. The World without Military Conflicts

is on the x-axis, while the y-axis gives the counterfactual one (the dashed line corresponds to the 45-degree line). Each dot is a dyad, and some are singled out by symbols: Grey diamonds represent intra-EU pairs; black crosses represent country pairs that were part of the communist bloc at some point; grey squares represent pairs that have had a nonzero frequency of recent wars in the real world.

Both EU and former communist country pairs experience a drop in their counterfactual probability of RTA formation with respect to the benchmark one. This is, we believe, another illustration of the window of opportunity channel. Indeed, for both groups of European countries the history of old wars is very intense. But due to the Cold War in particular, recent history was more peaceful as the two blocs were very stable internally between the end of WWII and the collapse of the USSR. Those 45 years of "forced" peace between countries that used to fight seems to have promoted the RTA wave in the region to a large extent.

Regarding the detrimental impact of recent wars, India-Pakistan and Egypt-Israel are probably the most illustrative examples. Those two pairs do have a very low level of benchmark probability of RTA formation, and this would jump to among the highest levels if one could cancel their history of recent wars. Greece-Turkey is another striking example.

Multilateralism Triggers Regionalism.—We now quantify the impact of multilateral trade openness on RTA formation. We simply cancel out multilateral globalization by setting multilateral trade openness to zero for all pairs of countries. We then estimate the resulting, counterfactual probability of RTA formation that we compare to the benchmark probability. Results are reported in Figure 7, where the grey triangles represent country pairs with an initial level of multilateral openness above the median level and where black circles represent pairs of countries belonging to Mercosur.

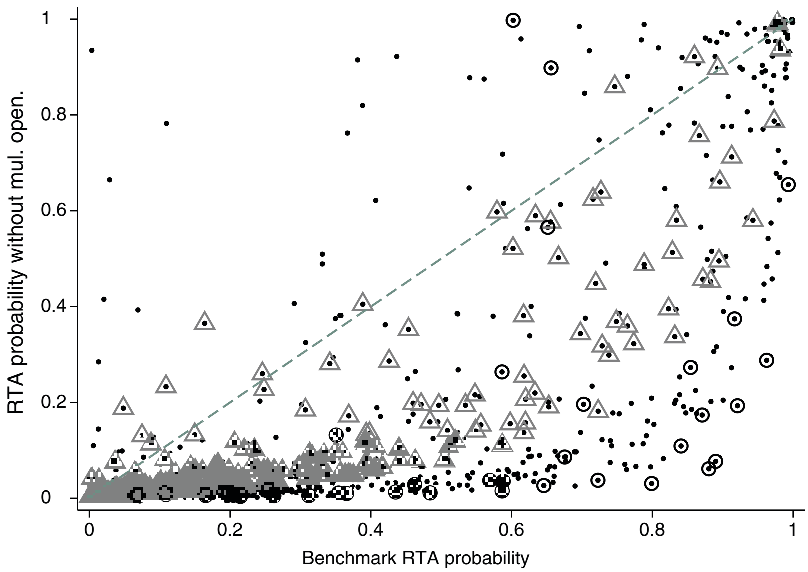

Figure 7. The World without Multilateral Trade

We see that in a counterfactual world without multilateral trade openness, most country pairs would experience a sharp decrease in their probability of RTA formation. This confirms our view that the wave of regionalism observed in the late 1990s could be a policy response to the destabilizing, and conflict-promoting, effect of the development of multilateralism as experienced during the post-World War period. This mechanism is particularly relevant for explaining the formation of Mercosur—a fact that has been discussed by policy practitioners (see Manzetti 1993).

IV. Conclusion

Our results suggest that political scientists and historians are right to emphasize the political motivation behind RTAs, in particular, the objective of pacifying relations. However, this does not mean that economics do not matter and that RTAs are signed without taking into account their economic benefits—trade gains. On the contrary, in absence of trade gains that may be lost during a war, the peace promoting effect of RTAs is greatly weakened. Hence, our story is one where politics and economics push in the same direction. Economic and security gains are complementary to explain the evolving geography of trade agreements. Trade gains may be instrumentalized for a superior objective of peace, but that makes them more, not less, important. Another important result is the interaction between multilateral and regional (or bilateral) trade liberalization. The recent multiplication of RTAs is often interpreted as a response of policymakers frustrated by stalling multilateral trade negotiations. Our result suggests a radically different story, one where multilateral openness (which may come from multilateral liberalization at WTO or the multiplication of RTAs) induces the formation of additional RTAs. RTAs can be interpreted as a way to reinforce bilateral economic relations between countries at risk of war at a time when globalization reduces the bilateral economic dependence of these countries. The domino theory of regionalism of Baldwin (1995) comes to mind. Here, the danger that additional RTAs are attempting to counter is not the loss of economic attractiveness but the dangerous loss of economic dependency that it may imply. Hence, RTAs may be contagious for political and not only for economic reasons. Finally, our results are consistent with the view that windows of opportunity for locking-in peace through trade exist. RTAs are difficult to sign for countries with a history of recent conflicts, but country pairs with a long-run history of bilateral conflicts have a higher propensity to sign an RTA. Hence, periods of peace between old enemies should be exploited to sign an RTA and lock-in a more peaceful bilateral relationship.

Appendix A: Theoretical Setup

In the absence of an RTA, the expected welfare of a country is equal to $V = (1 - \delta e)U_P + \delta e(1 - W)U_P$. Indeed, peace occurs with probability $(1 - \delta e)$ and, in that case, the country trades under MFN tariff with its partner and gets the benchmark welfare U_P. War occurs with probability (δe), and, in that case, trade is fully disrupted and some destructions happen; the country gets $(1 - W)U_P$. If an RTA is in force, the logic is similar and the expected welfare of the country is equal to $V^{RTA} = (1 - \delta e^{RTA})(1 + T)U_P + \delta e^{RTA}(1 - W)U_P$. Plugging those two expressions into the RTA formation condition, $V^{RTA} - V \geq C$, and rescaling by the benchmark welfare U_P, we easily obtain condition (2) in the main text.

Combining (2) and (3), we obtain

$$(1 - \delta e)T - \delta eW\left(1 + \frac{T}{W}\right)\frac{e^{RTA} - e}{e} \geq c \tag{A1}$$

$$(1 - \delta e)T + \delta eW\varepsilon_{cost}\left(1 + \frac{T}{W}\right)\left(\frac{\varepsilon_{pol}}{\varepsilon_{cost}} + \frac{T}{W}\right) \geq c. \tag{A2}$$

We assume that the RTA-related trade gains are small with respect to the cost of wars such that $T/W \sim 0$. Hence, we get the following approximation:

$$(1 - \delta e)T + \delta e(\varepsilon_{pol}W + \varepsilon_{cost}T) \geq c, \tag{A3}$$

which corresponds to equation (4) in the main text.

Appendix B: Marginal Effect and Interaction

The coefficient of the interaction term between trade gains and old wars is positive both in our benchmark LPM specification (column 7, Table 2) and in our benchmark logit specification (column 1, Table 4). However, Ai and Norton (2003) show that interaction terms have a sign that can be deceptive in a logit framework, and

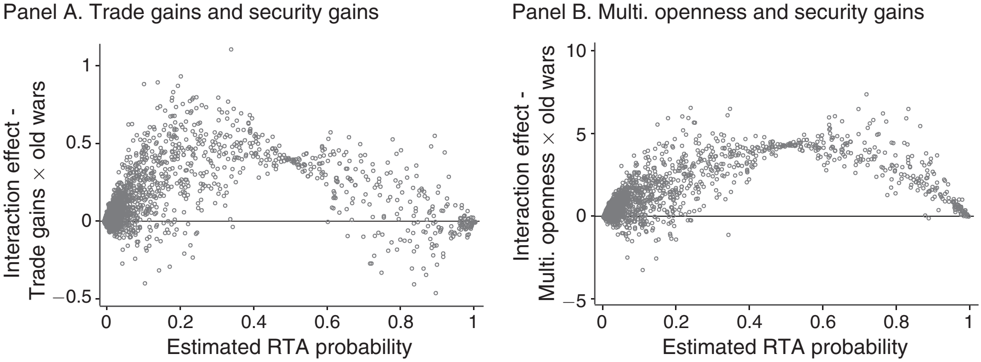

Figure A1. The Interaction Terms

that cannot be interpreted readily. Hereafter, we investigate this question more fully by calculating the marginal effect of this interaction term for the whole range of benchmark probabilities.

The interaction between old wars and trade gains, but also with multilateral openness, somewhat complicates the computation of the marginal effects with respect to Ai and Norton (2003). Let us denote x_1, x_2, x_3 our three variables of interest and $\mathbf{Z}$ the vector of covariates. Our logit preferred specification (8) writes as

$$\text{(A4)}\quad \hat{\mathbb{P}} = \frac{1}{1 + \exp[-\beta_1 x_1 - \beta_2 x_2 - \beta_3 x_3 - \beta_{12} x_1 x_2 - \beta_{13} x_1 x_3 - \beta \mathbf{Z}^t]},$$

where $\hat{\mathbb{P}}$ is the estimated probability of RTA formation. Simple computations lead to

$$\text{(A5)}\qquad \frac{\partial^2 \hat{\mathbb{P}}}{\partial x_1 \partial x_2} = \hat{\mathbb{P}}(1 - \hat{\mathbb{P}})\beta_{12} + \hat{\mathbb{P}}(1 - \hat{\mathbb{P}})(1 - 2\hat{\mathbb{P}})(\beta_2 + \beta_{12} x_1)$$
$$\times\ (\beta_1 + \beta_{12} x_2 + \beta_{13} x_3).$$

We use this formula to estimate the marginal effect of this interaction term for the whole range of benchmark probabilities. Results are shown in Figure A1. Panel A reports the marginal effects for the interaction between trade gains and security gains, while panel B reports them for the interaction between multilateral openness and security gains. In both graphs, each dot corresponds to an observed country pair. We see that the marginal effects of the two interaction terms are mostly positive. It also confirms that, due to the functional form of the logit probability distribution, the reversal of the sign of the marginal effects is more likely when the estimated probabilities are located in the neighborhoods of 0 and 1. Since, in our sample, those estimated probabilities are quite concentrated at those two extreme values, verifying that those marginal effects are indeed positive was important.

Appendix C: Further Country Pairs in Trade Gains Table

Table A1—Estimated Trade Gains for the 51st–100th country pairs in 1956

Country pair		Trade gains: Min T (percent)	Trade gains: Mean T (percent)	Bil. open. $\min \frac{imports}{GDP}$ (percent)	Dist. kms	Ever RTA?
USA	COL	0.259	2.262	0.105	4,251	No
FRA	**ITA**	**0.256**	**0.338**	**0.261**	**892**	**Yes**
THA	IDN	0.256	0.305	0.615	2,306	Yes
GBR	DEU	0.256	0.29	0.526	809	Yes
NLD	IDN	0.249	0.548	1.363	1,1346	No
CAN	VEN	0.248	0.262	0.683	4,647	No
BEL	SWE	0.244	0.3	0.941	1,152	Yes
DEU	DNK	0.239	0.845	0.425	538	Yes
CZS	BGR	0.236	0.72	0.307	1,084	No
GBR	ZAF	0.234	1.474	0.481	9,489	Yes
JPN	IDN	0.23	0.731	0.329	5,482	No
FRA	SWE	0.227	0.262	0.231	1,616	Yes
HUN	ROM	0.225	0.235	0.128	540	Yes
SAU	IND	0.223	0.3	0.191	3,509	No
DEU	AUT	0.222	1.358	0.393	592	Yes
CHN	LKA	0.219	0.426	0.095	4,914	No
CHN	JPN	0.216	0.3	0.167	1,975	No
FRA	CHE	0.214	0.57	0.217	474	Yes
ARG	GBR	0.213	0.3	0.438	11,137	No
CZS	ROM	0.212	0.456	0.275	902	No
HUN	BGR	0.211	0.256	0.278	693	Yes
GBR	IRL	0.209	2.164	0.429	425	Yes
BRA	SWE	0.209	0.247	0.545	10,185	No
IND	PAK	0.208	0.271	0.178	1,238	No
VEN	NLD	0.207	0.226	0.571	7,972	No
POL	AUT	0.203	0.211	0.227	549	Yes
BGR	ROM	0.2	0.257	0.105	370	Yes
BRA	URY	0.193	0.568	0.368	2,168	Yes
SDN	EGY	0.193	0.644	0.462	1,736	Yes
USA	BEL	0.192	0.601	0.078	7,303	No
BRA	DNK	0.19	0.191	0.365	9,776	No
ROM	EGY	0.19	0.198	0.1	1,792	No
POL	HUN	0.19	0.298	0.196	520	Yes
ARG	ITA	0.188	0.298	0.281	11,214	No
CHL	ARG	0.184	0.338	0.255	1,157	Yes
SYR	SAU	0.182	0.278	0.686	1,463	No
BRA	FIN	0.179	0.219	0.34	10,749	No
HUN	CHN	0.178	0.336	0.077	7,710	No
GBR	BEL	0.177	0.417	0.363	448	Yes
IDN	AUS	0.175	0.211	0.501	5,078	No
CHE	AUT	0.175	0.277	0.587	576	Yes
ARG	DEU	0.174	0.473	0.309	11,646	No
BRA	ESP	0.174	0.201	0.206	7,821	No
LBN	SAU	0.172	0.226	0.648	1,417	No
HND	SLV	0.171	0.274	0.519	244	Yes
JPN	AUS	0.171	0.404	0.346	7,827	No
SYR	JOR	0.169	0.317	0.733	373	No
BRA	NOR	0.168	0.169	0.324	10,018	No
AFG	PAK	0.168	0.257	0.104	806	No
ITA	SWE	0.166	0.194	0.248	1,833	Yes

Note: Lines in boldface indicate pairs that sign the Rome Treaty establishing the European Economic Community a year later.

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