Sovereign Gravity

The Military Alliance Effect on Trade*

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July 25, 2023

Abstract

International insecurity can severely disrupt trade. This paper studies treaties aimed at preventing such insecurity: military alliances. Taking a structural gravity approach based on a sample of 6,972 country pairs from 1967 to 2012, we show that alliances increase trade by 60% on average. Yet, the effects of military alliances are highly heterogeneous. They depend to a large extent on the type of alliance and the sensitivity of economies to trade costs. We use robustness tests and the instrumental variables strategy to confirm the causal interpretation of the results. Investigating the mechanism behind the impacts of military alliances, we demonstrate that alliances increase trade by reducing international insecurity. General equilibrium analysis moreover shows that the growth in trade generated by military alliances brings substantial welfare gains for signatories and losses for non-aligned countries.

Keywords: : Military Alliances, Trade, Structural Gravity, International Insecurity, Trade Elasticities, Welfare

JEL-Classification: F13, F15, F51, F62, D74

*We would like to thank seminar participants at the European University Study Group, Louis-André Gérard-Varet Days, International Atlantic Economic Society, Association Française de Sciences Economique, Eidgenössische Technische Hochschule Zürich, University Paris Dauphine, University Paris 1 Sorbonne, Paris School of Economics, Royal Holloway, Royal Economic society, International Trade and Finance Association, Mines Paris, Institutions Trade & Economic Development network, as well as James E. Anderson, Jan Bakker, Tibor Besedes, Lorenzo Caliendo, Natalie Chen, Mathieu Couttenier, Jose De Sousa, Jessica Di Salvatore, Peter Egger, Mario Larch, Emanuelle Lavallée, Thierry Mayer, Marion Mercier, Gianluca Orefice, Ralph Ossa, Mathieu Parenti, Ariell Reshef, Michele Ruta, Gianluca Santoni, Mike Spagat, Davide Suverato, Mathias Thoenig, Farid Toubal, Yoto Yotov.

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Since 2000, the number of major conflicts has quadrupled worldwide to involve more than 130 countries.[1] This massive increase in international insecurity can severely disrupt trade. The vast majority of trade costs are not associated with direct policy instruments, but with hidden transaction costs (Anderson and Van Wincoop, 2004). A significant proportion of these hidden transaction costs have to do with the insecurity of trade (Anderson and Marcouiller, 2002). Blomberg and Hess (2006)) show that common exposure to violence reduces bilateral trade by 7%, a figure that increases to 35% in the case of civil war. Sandkamp et al. (2022) determine that each additional maritime piracy incident reduces bilateral exports by 0.1%. Rohner et al. (2013); Yu et al. (2015) demonstrate the importance of conflict signals in shaping international trade costs. Martin et al. (2008) show that military interstate disputes reduce bilateral trade by 38% on average. Glick and Taylor (2010) demonstrate the overall negative impact of war on trade. They show that major interstate conflicts reduce trade by 80% between enemies and by 13% between belligerents and neutral countries, with a significant lasting effect in peacetime.

This paper examines one way of reducing trade insecurity: *the military alliances*. These international agreements are specifically designed to reduce insecurity among their members. They are based on two pillars: (1) enforcement of military cooperation policies, and (2) international security as a way to promote trade.[2] Many alliances exist such as the North Atlantic Treaty Organization (NATO), the Treaty on Collective Security and the Arab-Maghreb Union. However, the *Pax Mongolica* is an iconic example of such an alliance. In the 13th and 14th centuries, this set of treaties[3] ensured the security and development of trade in Eurasia. The end of the agreement saw a huge increase in conflicts and a sharp drop in trade – enough to prompt Europeans to take an unprecedented step in search of new trade roads (Findlay and O'Rourke, 2009). There are basically two categories of alliance: weak alliances – reducing the probability of open conflict between signatories – and defence pacts – enforcing collective and centralised management of members' security (Gibler, 2008). Their depth of military cooperation is significantly different, as is their expected effects on insecurity and trade.

Few previous papers have analysed the impacts of alliances on trade. Those that do are restricted to the Cold War period, find heterogeneous results and lack theoretical grounding. Drawing on the Tinbergen (1962) gravity equation, Mansfield and Bronson (1997) find a positive correlation between

[1] Authors calculation based on geocoded UCDP project data on conflict events (Sundberg and Melander, 2013).

[2] "An alliance is a formal contingent commitment by two or more states to some future action. The action involved could entail almost anything—detailed military planning, consultation during a crisis, or a promise by one state to abstain from an upcoming war. [...] empirical studies have developed a consensus that the operationalisation of the alliance variable depends on two factors. First, alliance members have to be independent nation-members of the international system (for example, so-called alliances between international terrorist organisations do not qualify), and second, a treaty text has to exist that identifies a military commitment that is defensive, a neutrality arrangement, or an "understanding" such as an entente," (Gibler, 2008). Details on alliances' content are provided in section A.

[3] The Pax Mongolica was a set of treaties between the former Mongol empire states – The Golden Horde (Western Steppe), the Yuan Empire (China), the Ilkhanat (Persia) and the Chagatai Khanate (Eastern Steppe) – the Italian republics and the Russian duchies (Findlay and O'Rourke, 2009).

exports and alliances using a panel regression covering the 1960-1990 period. They conclude that an alliance increases exports by 20%, a Regional Trade Agreement (RTA) by 49%, and both by 65%. Taking a similar approach, but with a generalised least squares estimator and no control for RTAs, Long (2003) estimates for the 1885-1990 period that defence pacts are associated with 37% higher exports, while weak alliances have no statistically significant effect.

Our study of the impact of military alliances on trade contributes to the literature in a number of ways. First, we identify the causal effect of alliances. On average, military alliances increase bilateral exports by 60%. Second, we investigate the mechanism behind this impact to show that alliances boost trade by significantly reducing international insecurity. Third, to the best of our knowledge, we are the first to investigate the welfare effect of the growth in trade induced by alliances. We show that the enforcement of an alliance brings substantial welfare gains for signatories, but losses for non-aligned countries.

This study is structured as follows. We present the focus on military alliances under structural gravity theory. We isolate the costs of insecurity in a model covering heterogeneous firms based on Chaney (2008) and Helpman et al. (2008). By reducing trade costs sensitive to insecurity, military alliances directly increase trade between partners. From this frame, we derive the gravity equation on which our empirical work is based. The analysis combines the Correlate of War database, built on the massive and meticulous work by Gibler (2008)) to document active military alliances, with the CEPII[4] CHELEM dataset on international trade. Thus, we perform a structural gravity approach using a panel of 6,972 country pairs covering the period from 1967 to 2012.

We estimate the effects of military alliances on bilateral exports. Taking exporter-year, importer-year and exporter-importer fixed effects, our specifications focus on the within-country-pair variation of military alliances. Our set of fixed effects ensures that we properly control for multilateral resistance terms, market access and structural interstate relationships (Behrens et al., 2012; Feenstra, 2015; Redding and Venables, 2004). In addition, we control for Regional Trade Agreements (RTAs). The spread of RTAs and alliances over our period is highly distinct. Yet, since we are investigating the specific effect of alliances, it is important to control for the standard agreements designed to affect trade. On average, enforcing a military alliance increases bilateral exports by 60%, which is equivalent to a tariff reduction of 12.8%.[5] This result is robust to a wide range of consistency checks, including additional controls (tariffs, depth of RTAs, Cold War, etc.), but also to other estimation techniques preventing potential bias (intranational trade, negative weights, asymptotic bias, etc.). Nonetheless, the effects of alliances are highly heterogeneous. They are sensitive to the nature of the treaty and the non-constancy of trade elasticities. Thus, only defence pacts have a significant effect, while small

[4]Centre d'Etudes Prospectuses et d'Informations Internationales.

[5]Equivalence is made with the estimated trade elasticity in our sample $\theta = 3.7$.

countries benefit more than large economies from the enforcement of such agreements.

We carefully investigate the endogeneity of military alliances. Using an instrumental variable strategy supported by a plausible exogeneity test based on Conley et al. (2012), and a Differenced Average Treatment on the Treated (DATT) analysis based on Couch and Placzek (2010), we confirm the causal interpretation of our results.

Then, we turn to analysing the mechanism by which military alliances affect bilateral trade. We directly test the validity and prevalence of the insecurity mechanism, i.e. the growth in trade driven by a reduction in insecurity. Retrieving data on conflict events from the geocoded UCDP project (Sundberg and Melander, 2013), we measure bilateral insecurity by interacting the country-time sums of conflict events, excluding military cooperative ones. Using a two-stage strategy and considering the heterogeneity of alliance treaties, we show that: (i) defence pacts sharply reduce bilateral insecurity, and (ii) by reducing insecurity, they significantly increase bilateral exports. The insecurity mechanism explains the effects of defence pacts on trade as a whole. Therefore, our results strongly support both the validity and prevalence of the insecurity mechanism.

In the last part of the paper, we investigate the welfare effect of alliances. We develop a general equilibrium analysis. In keeping with Arkolakis et al. (2012), we derive the welfare system from our theoretical model, pointing up the role of insecurity costs. Then, using the properties of the PPML estimator, we solve this system and perform a counterfactual analysis for 2012 in which all alliances are ended. This enables us to draw conclusions about the impacts of military alliances on real revenue – our measure of welfare. Military alliances bring their members substantial welfare gains. Interestingly, our results show that neutral countries experience a marked welfare cost at the same time. Moreover, performing a scenario analysis, we show the considerable potential welfare ramifications of reshaping the military alliance network in response to the war in Ukraine.

The paper is organised as follows. Section 1 describes the theoretical framework. Section 2 presents the data used in the analysis and some descriptive evidence. Section 3 investigates the effects of military alliances on bilateral exports, the sensitivity of our baseline results and heterogeneous effects. Section 4 addresses potential endogeneity concerns. Section 5 studies the mechanism through which alliances affect trade. Section 6 develops the general equilibrium analysis and draws conclusions about the welfare impact of military alliances. Lastly, section 7 presents a short conclusion.

1 Theoretical framework: Insecurity and structural gravity

When two countries sign an alliance, they enforce bilateral military cooperation policies (Gibler, 2008) – see appendix A for details and examples. In so doing, they improve security between partners.

This implies that when companies export from one signatory country to another, they have (i) less of an expropriation risk (destruction, political expropriation or robbery) and (ii) lower insecurity barriers (controls, procedures and information requirements). In other words, military cooperation is supposed to reduce insecurity costs (Anderson and Marcouiller, 2002). The expropriation risk is a variable cost while insecurity barriers induce fixed costs. This means that the potential effect of alliances is a reduction in both variable and fixed trade costs.[6] Below, we present the focus on military alliances under structural gravity theory and the resulting gravity equation. The full model is detailed in the appendices (Section B).

When a firm exports in the presence of insecurity, each product has a probability S_{ij} of being sold and $1 - S_{ij}$ of being expropriated. Firms export a number of units of their variety, so the probability of expropriation can be interpreted as the share of exports that does not arrive at destination (i.e. the consumer). This is similar to a traditional iceberg trade cost τ (Anderson and Marcouiller, 2002), but sensitive to insecurity with $\tau_{ij}(s) = 1/S_{ij}$.[7] In monopolistic competition with a Constant Elasticity of Substitution (CES) demand function, this leads to the price function:

$$p_{ij}(\alpha) = \frac{\sigma}{\sigma - 1} w_i T_{n,ij} \alpha \tag{1}$$

where σ is the elasticity of substitution, w_i the wage in country i, α the firm's marginal cost (i.e. the inverse of productivity γ). $T_{n,ij} = \Pi^n \tau_{n,ij}$ is a product of variable trade costs with n the n potential source of iceberg cost, including $\tau_{s,ij}$ the variable insecurity cost (derived from the expropriation risk), but also all variable trade costs sensitive to other parameters (geography, standard trade policies, institutions, etc.). Thus, in such a frame, any reduction in bilateral insecurity reduces exporting firms' prices.

Second, we introduce insecurity barriers. Firms need to address expensive procedures and information requirements to lift these barriers and enter the foreign market. The higher the bilateral insecurity, the higher the barriers and therefore the higher the cost to lift them. This is directly interpretable as a fixed trade cost sensitive to insecurity: to enter the market, firms pay a cost that depends on the degree of bilateral insecurity but does not vary with the exported quantity. So it does not affect the

[6] We do not make any composition assumptions. Military cooperation policies described in appendix A suggest that both variable and fixed insecurity costs are reduced by alliances. In our analysis of the overall effect of alliances and the role of bilateral insecurity, we recognise for interpretation reasons the existence of both variables in fixed insecurity costs, but do not need to disentangle them.

[7] We can also interpret the reduction of the expropriation risk as a lower cost of insurance. In the case of an insurance market, a firm can pay insurance which, in exchange for a contribution equal to the share $\tau_{In,ij}$ of the value of each insured exported product, will provide the amount p (the price) for each expropriated product. Thus, in exchange for a variable cost $\tau_{In,ij}$, the firm obtains the guarantee that the exported products will be sold at the price p. Given that we are in a Melitz (2003) monopolistic competition case, firms face a returns-to-scale technology due to the presence of fixed costs and are price setters. Therefore the firm chooses the lower price between $p_{ij}(\alpha) = \frac{\sigma}{\sigma-1} w_i \tau_{s,ij} \alpha$ and $p_{ij}(\alpha) = \frac{\sigma}{\sigma-1} w_i \tau_{In,ij} \alpha$.Firms only take out insurance policies that will ensure $\tau_{In,ij} <= \tau_{s,ij}$. So, when the probability of expropriation decreases, there is a reduction in $\tau_{In,ij}$. Even in the presence of an insurance market, military alliances reduce bilateral variable trade costs.

price function, but directly affects profit:

$$\pi_{ij} = (\frac{x_{ij}(\alpha)}{\sigma}) - F_{n,ij} \tag{2}$$

$x_{ij}(\alpha)$ is the firm's revenue function and $F_{n,ij}$ a vector of fixed trade costs that firms have to pay to enter country j from i, including $f_{s,ij}$, the fixed insecurity cost derived from insecurity barriers, but also all fixed trade costs sensitive to other sources n. If the market entry cost decreases, the firm's profit increases. When firms switch from negative profit to positive profit, they start exporting, which increases the number of varieties sold from i to j.

From this theoretical frame, we can derive at aggregated level the following structural gravity equation outlying the insecurity trade costs $\tau_{s,ij}$ and $f_{s,ij}$:

$$X_{ij} = N_i \overline{\alpha}_i^{-\theta} w_i^{-\theta} \frac{X_j}{\Phi_j^{-\theta}} \tau_{s,ij}^{-\theta} f_{s,ij}^{-[\frac{\theta}{\sigma-1}-1]} T_{n\neq s,ij}^{-\theta} F_{n\neq s,ij}^{-[\frac{\theta}{\sigma-1}-1]} \tag{3}$$

X_{ij} is the total exports from country i to country j, N_i, the number of firms in the exporting country, $\overline{\alpha}_i$, the maximum marginal cost of country i's technology, and w_i the wage in country i's economy. X_j is the total revenue of country j, and Φ_j, the importer's multilateral resistance term, while $T_{n\neq s,ij}$ and $F_{n\neq s,ij}$ are respectively the variable and fixed trade costs sets, but excluding the insecurity costs. Like other trade costs, insecurity costs ($\tau_{s,ij}$ and $f_{s,ij}$) have negative elasticities.[8] Hence, by reducing the insecurity costs, the enforcement of a military alliance between countries i and j increases bilateral exports X_{ij}.

2 Data

The structure of the dataset is a country-pair panel. Our unit of observation, therefore, is a given exporter-importer-year combination. We study how variations in the ally status of the dyad affect bilateral exports.

2.1 Data description

Alliances data.– We use information on military alliances for each ijt from the Correlate of War project (Gibler, 2008). We have information on whether a given country pair are allies and, if so, the nature of the treaty. We can divide military alliances into two categories: weak alliances, which

[8] These elasticities depend both on θ, the Pareto shape parameter of the firms' productivity distribution. Yet, fixed trade cost elasticity also depends on σ, the elasticity of substitution. Here, we assume $\theta > \sigma - 1$. Otherwise, fixed trade costs elasticity is positive. Note also that despite our assumption that all trade costs of the same nature (variable vs fixed) have the same elasticity, this does not mean that the model assumes that all policies have the same trade elasticity. Indeed, we do not assume that all trade costs are sensitive to the same policies or to the same extent. Details are provided in section B.

focus mainly on military cooperation to guarantee peace between signatories, and defence pacts, which enforce military cooperation to protect members from outside threats and achieve common strategic objectives[9] (Long, 2003; Gibler, 2008). From 1967 to 2012, the majority of military alliances were defence pacts. Yet, most of them were enforced throughout the entire period. Consequently, in our sample the number of defence pacts and weak alliances contributing to the within variation – i.e. whose status changes over the period – is comparable (see table 1). For each country pair, we define ALL_{ijt}, a dummy variable which equals 1 if country i and j are allies at time t and 0 otherwise.

Insecurity and military cooperation data.– We collect data on conflict events from the geocoded UCDP project (Sundberg and Melander, 2013) to construct our measures of military cooperation and bilateral insecurity. The initial observation unit is an event. Information is available starting in 1989 with the year provided for each event. This project also has the advantage of identifying the belligerents (and co-belligerents) in each conflict event. We organise the information to create a dummy taking the value one if the country-pair cooperates militarily (i.e. is belligerent) in the event. Summing this dummy at country-pair-year level, we obtain a continuous measure of bilateral military cooperation. By summing observations of conflict events, excluding the country-pair's cooperative events, we observe country-year exposure to insecurity.

Exports data.– International trade data are retrieved from the CEPII CHELEM base (de Saint-Vaulry, 2008). We extract bilateral exports in current dollars between the 84 available countries from 1967 to 2012.[10] Data exclude re-imports and re-exports. Flows are adjusted for freight and insurance costs.[11] Zero trade flows are observed. Our export matrix is squared. The CHELEM database enables us to exclude arms exports, which improves our identification strategy. Therefore, in our case, CHELEM provides the best trade-off between quality of observations and panel size. Nevertheless, we discuss below a robustness check with extended trade data. The CHELEM base does not provide information on intranational trade. We supplement our export data with within-country flows from the CEPII TradeProd database (De Sousa et al., 2012; Mayer et al., 2023).[12] Similarly to CHELEM, flows are adjusted for freight and insurance costs and harmonised according to the reliability of the countries' declarations. However, the dataset starts in 1990 and 19 CHELEM sample countries are unobserved.[13] Within-country trade flows are very important for the general equilibrium analysis

[9] See section A for more information about military alliance treaties and examples.

[10] The full list of countries is presented in appendix table 11.

[11] In declarations, imports include freight and insurance costs while exports do not. Considering the reliability of countries' declarations, CHELEM's bilateral trade is harmonised in keeping with the RAS iterative method (see Stone (1963)). Prior to 1992-93, some countries, such as the former USSR and former Yugoslavia, are not recognised (or reported) as independent trade partners by the UN. The CHELEM base provides estimated values to fill these missing observations. The dataset therefore contains harmonised export values for all exporter-importer pairs (6,972) across the entire period.

[12] Intranational flows are filled by linear interpolation of non-missing data, whereas the remaining missing values are extrapolated using country total exports (Baier et al., 2019; Fontagné and Santoni, 2021).

[13] Excluded countries are: Belarus, Bosnia and Herzegovina, Brunei, Croatia, Czech Republic, Estonia, Kazakhstan, Kyrgyzstan, Latvia, Libya, Lithuania, Luxembourg, North Macedonia, Paraguay, Russia, Slovakia, Slovenia, Ukraine.

performed in the last part of the paper. Yet, because of this data limitation, the baseline and the other partial equilibrium estimations use only international exports. In this way, we make use of the largest available panel. Nonetheless, a robustness check with intranational flows is discussed below.

Other data.– Information on Regional Trade Agreements (RTAs) and the standard gravity variable, such as distance, population, common language and religion, are retrieved from the CEPII's Gravity database (Head et al., 2010).). Data on RTAs include preferential trade agreements, free trade agreements, customs unions and other less common forms of agreements. We round them out with RTA legally enforceable provisions from the Content of Deep Trade Agreements database (Hofmann et al., 2017).). Information on Gross Domestic Product (GDP) is also taken from the CHELEM dataset. Information on tariffs is retrieved from the World Trade Integrated Solution, which combines data from UNCTAD TRAINS[14] and the World Trade Organization. Finally, we extract our data on corruption and the rule of law from the Variety of Democracy project (Coppedge et al., 2022; Pemstein et al., 2022).

2.2 Descriptive statistics

We have a panel of 6,972 country-pairs from 1967 to 2012. In our dataset, 46% of worldwide exports (in value) come under the umbrella of alliances: 72 countries are signatories to these treaties and 740 pairs are affected by one, including 364 making a switch during our period (cf. table 1).

In figure 1, the map of the world displays the number of alliances per country during our time-frame. As can be observed, alliances are heterogeneously distributed across countries. No clear correlation between level of economic development and being signatories to such treaties is observed. Intermediary or low-income countries are not excluded from the worldwide alliance system – South American countries have signed more alliances than any European countries, while Africa and Asia present a wide range of involvement. In figure 2, the same exercise is replicated with the number of switches in alliance per country. Countries contributing to the switches are well dispersed around the globe, providing a good range of treated economies and international relationships.

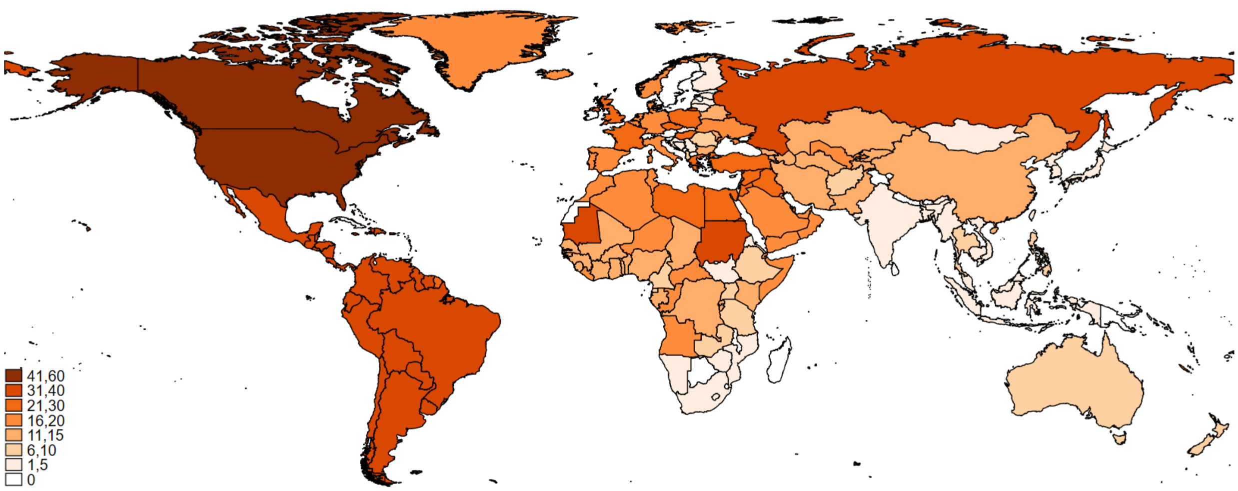
[14] United Nations Conference on Trade and Development Trade Analysis Information System.

Table 1: Alliances and RTAs, countries involved

	Alliances	Defence pacts	Weak alliances
Countries involved	72	63	45
Countries never involved	12	21	38
Country-pairs involved	738	618	142
Country-pairs with a switch	362	244	142

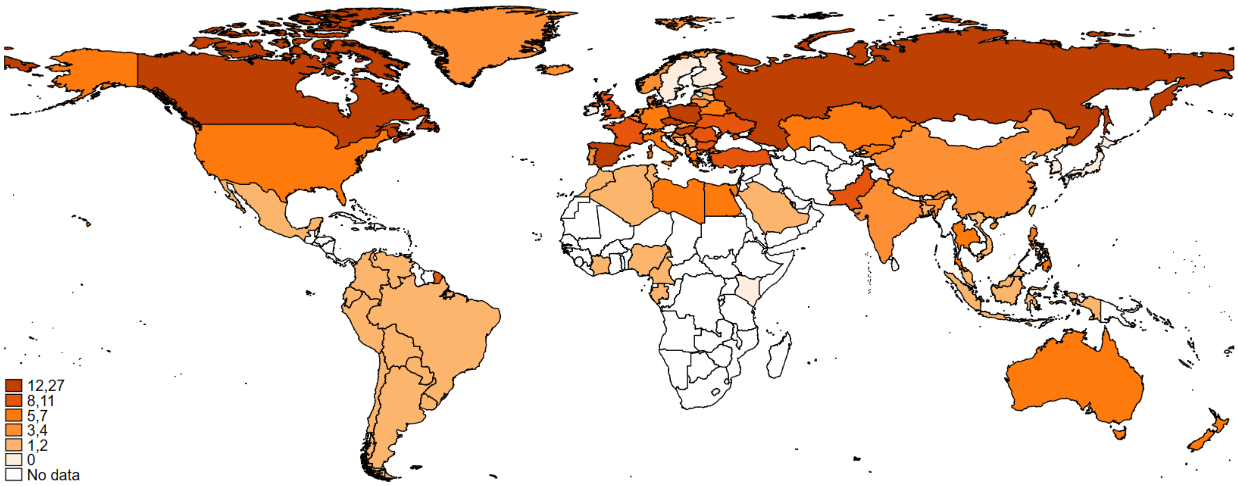
Note: Author's calculation. We count each exporter-importer observation as a country-pair. All alliances are symmetric.

Figure 1: Number of alliances by country, 1967-2012 (COW data)



Note: alliances are counted at country-pair level; 60 means that the country has been allied with 60 other countries from 1967 to 2012; white areas are where no alliance has been observed.

Figure 2: Sample switches in alliance by country

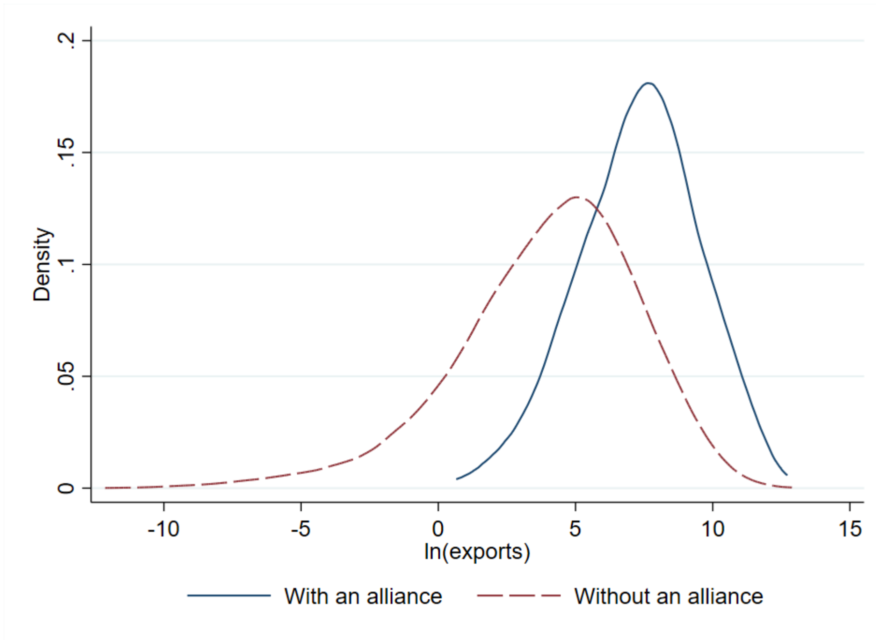


Note: Alliance switches are counted at country-pair level; 27 means that the country has signed or terminated a military alliance with 27 other countries from 1967 to 2012; white areas are countries excluded from our final sample.

A simple density graph (cf. figure 3) displays a positive correlation between bilateral exports and military alliances. The distribution of country-pair exports with military alliances lies more to the right than the distribution without, indicating a significantly higher average level of exports for country pairs with a military alliance than without.

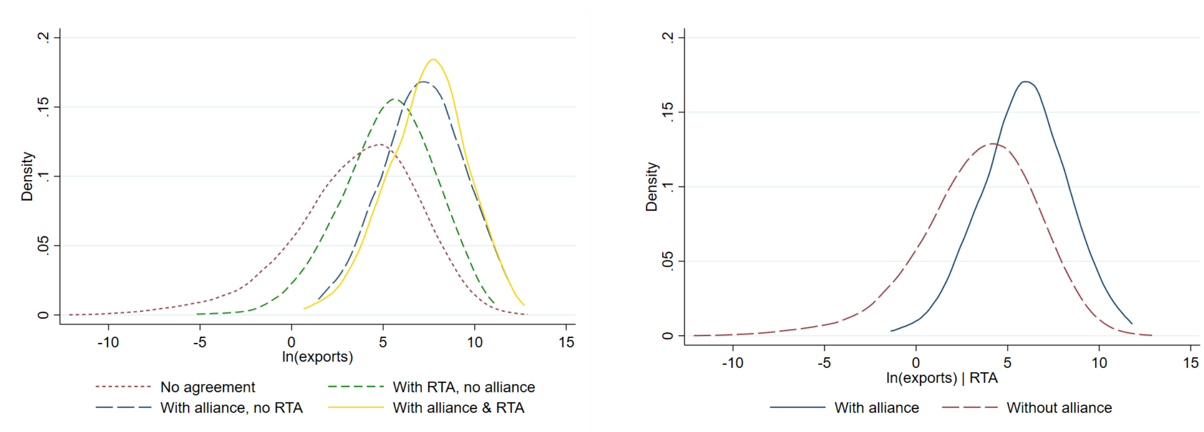
RTAs are supposed to be international trade liberalisation agreements, but the correlation between alliance enforcement and exports does not seem to be very dependent on the existence of RTAs. Figure 4 presents a further two density graphs. Graph 4a, displays export values depending on whether the pair has an RTA, an alliance, both or no agreement. Graph 4b reproduces graph 3 but with export values conditional on RTAs (i.e. exports unexplained by RTAs). Both graphs present interesting evidence that, irrespective of the existence of an RTA, enforcement of an alliance is positively correlated with bilateral exports.

Figure 3: Export values and alliances



Note: The K-density graph compiles exporter-importer exports for the latest year of available data (2012).

Figure 4: Export values, alliances and RTAs



(a) Export values

(b) Export values conditional on RTAs

Note: K-density graphs compile exporter-importer exports for latest year of available data (2012).

2.2.1 The role of alliances in insecurity

Table 2, regresses standard proxies of insecurity on ALL_{ijt} and RTA_{ijt}. These insecurity proxies are drawn directly from the literature presented in the introduction. First, we replicate the approach taken by Anderson and Marcouiller (2002) using an institutional proxy designed to capture the level of corruption. As in Anderson and Marcouiller (2002), institutional changes are observed at country-level and measured in log-ratios.[15] In addition, country i and j corruption log-ratios are interacted to obtain bilateral changes. Second, in keeping with Rohner et al. (2013); Yu et al. (2015), we derive insecurity costs from conflict signals, which are observed in terms of institutional law enforcement differences. Thus, in column (2), we use as a proxy for insecurity the log of the absolute-value difference in rule of law between countries i and j. Third, as in Blomberg and Hess (2006) we create a conflict dummy taking the value 1 if country i and j are both exposed to at least one conflict event at time t and 0 otherwise.[16] In this case, bilateral insecurity is measured using the extensive margin of country conflict exposure. The finding is the same in each column of table 2: the coefficient of ALL_{ijt} is statistically significant, negative and much larger than for RTA_{ijt} (see F-tests).[17] In other words, alliances are closely correlated with a lower level of bilateral insecurity, while RTAs are not.[18]

Table 2: Descriptive correlations: bilateral insecurity, alliances and RTAs

	(1)	(2)	(3)
Dependent variable:	Corruption	Rule of law(diff.)	Conflict dummy
Alliance	-0.278^a	-0.066^a	-0.088^a
	(0.050)	(0.007)	(0.013)
RTA	-0.003	-0.037^a	-0.013^b
	(0.037)	(0.003)	(0.014)
Controls	yes	yes	yes
Country i x Year FE	no	yes	yes
Country j x Year FE	no	yes	yes
Dyadic FE	no	yes	yes
No. observ.	163,484	262,324	167,328
All-RTA F-test	18.38	26.63	27.60

Note: The estimator is Ordinary Least Squares; Dependent variables are different proxies for bilateral insecurity as detailed in section 2. Robust standard errors clustered at country-pair level are in parentheses.
Controls are $ln(GDP_{it})$, $ln(GDP_{jt})$, $ln(pop_{it})$, $ln(pop_{jt})$, $ln(distance_{ij})$, $common.religion_{ij}$, $common.language_{ij}$, $colonial.past_{ij}$, $contiguity_{ij}$; these are all captured by fixed effects in estimations (2) and (3). Difference in observation numbers are due to data availability restrictions. a, b and c denote significantly different from 0 at the 1%, 5% and 10% level, respectively..

[15] Formally: $ln(\frac{corruption_{it}+1}{E_{W t} corruption_{it}+1})$ and $ln(\frac{corruption_{jt}+1}{E_{W t} corruption_{it}+1})$, where W means the world.

[16] Contrary to Blomberg and Hess (2006), the Conflict dummy used here is not restricted to non-state events. Thus, the variable also takes the value one in the case of violent events involving state forces.

[17] Note that we do not use the obvious War dummy measure as a proxy. Unlike Glick and Taylor (2010), our modern timeframe means that our panel contains very few interstate war observations. Moreover, the correlation between alliances (especially defence pacts) and war is very close to -1, since almost no country-pair with a switch in ALL_{ijt} was involved in an open war with each other. Therefore, we cannot regress a War dummy on alliances. Note here also that modern insecurity takes neither exclusively nor mainly the form of open interstate wars, but also exposure to (potential) violence from states or organised non-state actors for whatever reason (geo-strategic, economic, ideological, etc.).

[18] Given that, as seen from our theoretical discussion, insecurity can lead to market entry barriers imposed by states, it is not surprising to observe that RTAs are not completely uncorrelated with insecurity variables since these treaties (especially the deepest) can include some agreements on barriers that affect the fixed costs.

3 Identification strategy

Following our theoretical discussion, the relationship between military alliances and exports can be estimated using a structural gravity model. Accordingly, our baseline specification is as follows:

$$X_{ijt} = exp(\beta_1 ALL_{ijt} + \beta_3 RTA_{ijt} + \lambda_{it} + \lambda_{jt} + \lambda_{ij}) * \epsilon_{ijt} \tag{4}$$

Our interest variable ALL_{ijt} is a dummy taking the value one if there is an alliance between country i and j at time t and zero otherwise. RTA_{ijt} is coded the same way as alliances, but for regional trade agreements. Alliances and RTAs may exist concurrently. Hence, we need to control for RTAs to capture any specific trade agreement effect between i and j.

λ_{it} and λ_{jt} the exporter-year and importer-year fixed effects. They capture the country, year and country-year-specific variables such as economic size and multilateral resistance terms (Baier and Bergstrand, 2007; Feenstra, 2015; Redding and Venables, 2004). We also include exporter-importer fixed effects (λ_{ij}) to capture any omitted variables due to structural relations between countries such as distance, common language and colonial past. Military alliances active throughout the period are also captured. We hence estimate the within-effects of military alliances (i.e. country pair changes in status).[19]

Military alliances are expected to impact on all sectors by reducing insecurity costs. Yet, military alliances can also be associated with arms supply contracts. We therefore exclude the arms sector from the bilateral exports variable X_{ijt} to make sure that what is measured is a trade cost reduction and not a contract effect.

The reverse causality argument is unlikely to bias our baseline estimation. An alliance is not an economic treaty, but a long-lasting military pact with heavy political constraints. Therefore, to find a pair-specific export shock affecting the signature of a military alliance is a remote possibility. Nonetheless, we test alternative specifications and address residual endogeneity concerns in a further section.

3.1 Baseline Results

Table 3 reports the baseline results. The dependent variable is the exports from country i to country j in year t. The effect of military alliances on bilateral exports is positive and significant. Enforcing a military alliance increases bilateral exports by 60% on average. By contrast, the average

[19] We use a Poisson Pseudo Maximum Likelihood (PPML) estimator to retain a non-linear specification and address heteroscedasticity. In this way, we take into account zero trade observations and avoid the biases caused by a combination of log-linearisation and heteroscedasticity (Silva and Tenreyro, 2006). In addition, standard-errors are clustered at exporter-importer level.

effect of RTAs is 17%. Translating the effects of alliances into tariff-equivalent variations under the standard trade elasticity calibration of $\theta = 3.7$[20] returns a tariff reduction of 12.8%. This equivalence appears reasonable given that we are dealing with a treatment assumed to have a sizeable impact on bilateral insecurity.[21] such equivalence appears reasonable. Nevertheless, in subsequent sections, we are careful to test the robustness of this result and focus on understanding the heterogeneity and mechanism behind it.

Table 3: Exports and military alliances

Estimator:	PPML
Dependent variable:	Bilateral exports
Variables	(1)
Alliance	0.473^a
	(0.106)
RTA	0.161^a
	(0.033)
Exporter x Year FE	yes
Importer x Year FE	yes
Dyadic FE	yes
No. observ.	320.666

Note: PPML, Poisson Pseudo Maximum Likelihood; FE, Fixed effects; Dependent variable is exports from country i to country j at time t in millions of current dollars. Standard errors clustered at country-pair level are in parentheses. a, b and c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

3.2 Sensitivity analysis

This section proposes a battery of robustness checks for our baseline results.

Intra-national trade flows.– Data limitations are such that using intranational trade flows would significantly reduce our sample and limit the observation of military alliances' within-variations (see section 2 on data for more details). Thus, the baseline estimation considers international trade flows only. However, this means that the alliance coefficient cannot include the potential effect of alliances in terms of diverting trade from intranational flows to international flows (Dai et al., 2014). Therefore, as a robustness check, we replicate the baseline estimation using a shorter panel containing within-country trade flow data. Results are displayed in 12 column 1.

Tariffs– We investigate the robustness of our results to the inclusion of tariffs. The tariff variable

[20]We directly estimate trade elasticity by including tariffs in the standard structural gravity estimation – see appendix table 23. The coefficient of $ln(1+tariffs)_{ijt}$ is directly interpretable as the trade elasticity(Anderson et al., 2018; Head and Mayer, 2014)

[21]This assumption is directly treated (and confirmed) in section 5

is defined as the log of the ad valorem exporter-importer-time average tariff rate plus one. The results of this sensitivity test are reported in table 12 column 2 to 4. The tariff coefficient is -0.492, but is not significant because of the poor country-pair-within variation of tariffs in the data. This issue is addressed by including within-country trade flows. The estimated tariff coefficient is now large and highly significant – see column 5. This additional control variable does not reveal any omitted variable bias. The lower coefficient of the alliance variable is merely the outcome of the considerable reduction in the timeframe and the country-pairs sample for our panel dictated by tariff data availability. In addition, controlling for common GATT membership has no effect.[22]

RTA depth– In the baseline estimation, dummy variable RTA_{ijt} for the presence of an RTA between exporter and importer. This ensures that the coefficient of ALL_{ijt} is not biased by concomitant variations in RTAs and alliances. Yet, this method does not control for changes in RTA depth. To address this point, we proxy RTA depth by the number of provisions in each agreement. We then introduce this new variable into our baseline estimation. We drop country-pair-year observations where an RTA is observed, but not its depth. Results are displayed in table 12 column 6-7. We keep the RTA_{ijt} dummy in column one. Its coefficient is not empirically interpretable since it corresponds to a fictive empty RTA (with no provision). In the second column, RTA_{ijt} is dropped. In both cases, the coefficient of ALL_{ijt} is barely affected.

Distance and economic development.– Geographic distance and long-term differences in economic development are already captured by the country-pair fixed effect. Yet, both bilateral exports and signatures of alliance treaties could be affected by regionalisation or globalisation, i.e. by variations in transport costs and differences in economic development over time. Therefore, we build $ln(distance_{ij}) * year_t$ as an interaction variable between the distance and the year [23], and $ln(|\frac{GDP_{it}}{Pop_{it}} - \frac{GDP_{jt}}{Pop_{jt}}|)$ as the log of the difference in per capita GDP in absolute value between exporter and importer. We include these variables in the baseline estimation. Results are reported in table 12 column 8. Their coefficients are weak and not significant, while the military alliance coefficient remains unchanged.

Extended Panel.– For quality reasons, our trade data limit our panel's sample of country-pairs and time-frame. We test the sensitivity of our results to alternative trade data allowing for a larger panel. We use the IMF DOTS database to extend our panel from 46 to 65 years, and from 6,972 to 36,868 country pairs. Yet there are costs involved. First, we no longer observe zero trade, which implies we can only capture the intensive margin.[24] Second, we do not observe all country-pairs over the entire

[22] We introduced a dummy variable taking value one if both country i and j are members of the GATT. The alliance coefficient and standard error are unaffected by this additional control.

[23] For ease of coefficient interpretation, the year variable is equal to the year minus 1966.

[24] i.e. the effects of alliances on countries that were already trade partners. See section 3.4 below for a discussion of extensive and intensive margins.

period. Third, data are not subject to the same harmonisation and verification process as CHELEM. This undermines quality, especially for developing countries. Fourth, we cannot exclude the arms sector.[25] Results are reported in table 12 column 9. We still observe a clear positive and significant effect of alliances on trade. Nonetheless, the DOTS data limitation and the change in panel mean that we cannot make a direct comparison with the baseline.

Asymptotic bias.– Table 13 presents the corrected FE-PPML estimation developed by Weidner and Zylkin (2021). The military alliance coefficient is affected by a small negative bias of the order of -0.01, while the associated standard error is slightly underestimated. Once the correction is applied, the alliance coefficient remains highly significant and very similar to the baseline.

Negative weights.– The baseline strategy is similar to a fixed-effects difference-in-differences estimation. Yet, the effects of military alliances could be dynamic. In this case, a different effect would be found depending on the duration of the treaty, where the baseline coefficient would be the average. Moreover, the effects of military alliances on trade could also be heterogeneous across time and country pairs. Therefore, our results could suffer from negative-weight biases (De Chaisemartin and d'Haultfoeuille, 2020). Thus, to address these econometric considerations, we regress equation (4) using the De Chaisemartin and D'Haultfoeuille (2022) estimator. The estimation is detailed in appendix section C. It concludes that signing an alliance increases bilateral exports by 40% at time t (the date of signature).The effect then gradually grows over time to attain 98% in $t + 5$. Since the baseline results are the average of the effects of military alliances over time, the estimated effect of 60% is consistent with the observed dynamic and robust to the negative weights bias.

3.3 Heterogeneity of alliance treaties

There can be two types of alliances: *the weak alliances*[26] and *the defence pacts*. This distinction is based on the particularity of defence pact military cooperation policies (Gibler, 2008). In a nutshell, defence pacts are the only alliances designed to protect members from the rest of the world (see appendix section A for more details). We do not consider this distinction in the baseline. Yet, because of this fundamental treaty difference, we might expect defence pacts to have more of an effect on trade than weak alliances. Long (2003) provides a test of this difference using a non-structural gravity approach, concluding that only defence pacts are positively correlated with trade. In this section, we investigate whether the difference in alliance treaty categories produces heterogeneous effects, estimating within-effects using a structural gravity model.

[25] Some arms exports information is provided, albeit less than by CHELEM. Correcting DOTS trade flows for arms exports would drastically reduce the panel size.

[26] Non-aggression pacts, neutrality pacts and ententes.

We decompose the alliance variable into: i) a dummy taking the value one if there is an alliance with a defence pact between i and j at time t, and ii) and ii) another dummy taking the value one if there is an alliance without a defence pact between i and j at time t. Then, we replicate our baseline estimation. The results are reported in appendix table 15 column 1. Defence pacts increase bilateral exports by 100%, while the weak alliance coefficient is not statistically significant. This shows that a high level of military cooperation policies is required for alliances to affect trade. Moreover, the fact that defence pacts account for the majority of alliances explains the intensity of our baseline result – the latter is driven by the effect of very deep agreements. Conclusions are robust to the Cold War and the particularities of international relations during this period. A full discussion about the Cold War is provided in appendix section D. In addition, we test whether our results are driven by the largest defence pact in our sample – NATO. We create a specific dummy variable for the treaty and retain the defence pact dummy variable for the others.[27] The estimation results presented in table 15 column 2 confirm our conclusions.[28]

3.4 Intensive and Extensive margins of trade

In this section, we investigate through which margin of trade military alliances affect bilateral exports. We first estimate the effects of military alliances on bilateral exports conditional on positive flows from country i to j at year t and $t-1$ (the intensive margin). Then, using a non-linear probability model (Kitazawa, 2012; Silva and Kemp, 2016), we estimate the effects of alliances on the probability of starting to export to a destination (the extensive margin).[29] Results are reported in appendix table 14. The intensive margin estimation results are reported in panel A of the table, while the extensive margin results are shown in panel B. First, alliances increase by 47% the bilateral exports of country-pairs that were already trade partners. Second, they increase by 35% country i's probability of starting to export to country j. Thus, military alliances affect bilateral exports in terms of both margins. This suggests the presence of heterogeneous effects depending on which margin applies to the country-pair.

3.5 Non-constant trade elasticities

In line with standard gravity theory, trade cost elasticities are constant in our model (Head and Mayer, 2014).[30] Yet, recent literature has shown that the distance effect is decreasing with the size of bilateral exports (Carrère et al., 2020). As developed in section 1, military alliances are expected to

[27] Thus, in this estimation, the defence pact variable takes the value 0 if $NATO = 1$.

[28] In keeping with the definition of military alliances given by Gibler (2008), some recent NATO members (post-2003) are considered to belong to the international organisation, but not the alliance. We test the validity of our estimation including all NATO members in the alliance system. Results are barely affected.

[29] Given that the Logit estimator does not allow for intensive use of fixed effects, we use only country-pair and year fixed effects, which partially capture the multilateral resistance while controlling for importer/exporter GDP and population.

[30] In equation 3, trade elasticities are determined by θ, the Pareto shape parameter of the productivity distribution, and σ the elasticity of substitution. Both parameters are assumed to be constant across countries.

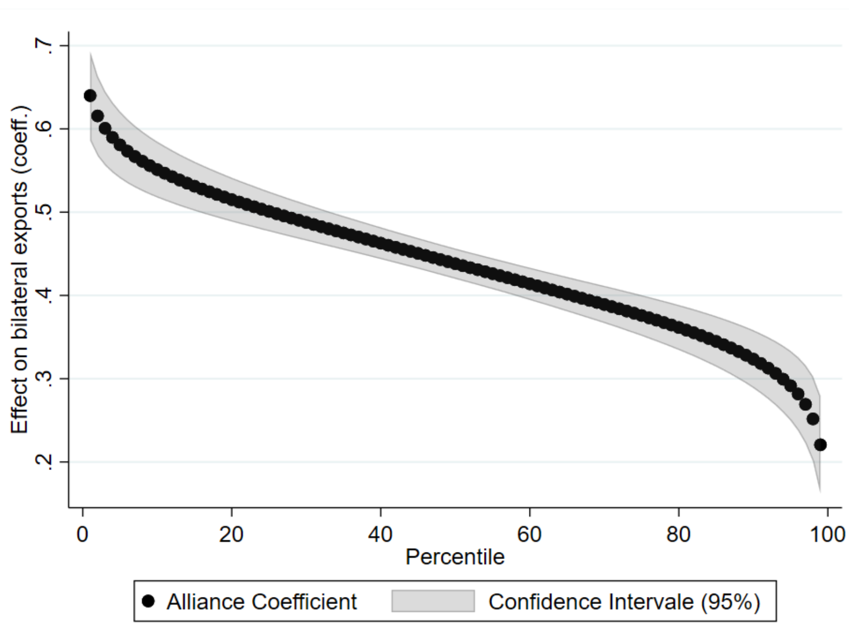
affect trade by reducing trade costs. Therefore, this section investigates whether the effects of military alliances on bilateral exports are concerned by non-constant trade elasticities.

In keeping with Carrère et al. (2020), we use the Method of Moments-Quantile Regression estimation (Machado and Silva, 2019):

$$\widetilde{X}_{q,ijt} = \beta_1 ALL_{q,ijt} + \beta_2 RTA_{q,ijt} + BZ_{q,ijt} + \lambda_{q,it} + \lambda_{q,jt} + \epsilon_{q,ijt} \tag{5}$$

$\widetilde{X}_{q,ijt}$ is the inverse hyperbolic sine transformation of $X_{q,ijt}$. $\lambda_{q,it}$ and $\lambda_{q,jt}$ are exporter-year and importer-year fixed effects, while q is a defined quantile and Z a vector of bilateral variables: distance, interaction between distance and year, colonial past, common language, common currency, common religion and territorial contiguity. Yet, given the sub-divisions performed by the quantile estimator, introducing dyadic-FE would entirely capture the effect of our variable of interest. The estimator is linear. So, the dependent variable is the natural log of exports augmented by one. We regress the equation for each percentile. The obtained $ALL_{q,ijt}$ coefficients, with bootstrapped 95% confidence intervals, are graphically displayed in figure 5. Similarly to the conclusions drawn by Carrère et al. (2020) with respect to distance, the alliance coefficients are decreasing in the value of trade. We plot the RTA coefficients in appendix graph 7. They also show a decrease in coefficients with the size of exports.[31]

Figure 5: Quantile estimates of alliances



[31] Another way to test the presence of non-constant trade elasticities is to change the weight of observations in our baseline. Using import share as the dependent variable, we increase the weight of observations with small export flows (Sotelo, 2019). We report the results in table 12 column 10. In line with our findings, we observe higher coefficients for both alliances and RTAs.

Two main conclusions are drawn from these results. First, the effects of military alliances on trade depend on the size of bilateral flows. In structural gravity, this implies that military alliances have heterogeneous effects depending on the size of the importer and exporter economies. Second, military alliance coefficients behave in the same way as variables that clearly affect trade through trade costs. This points to an alliance effect through a trade cost reduction mechanism. In appendix section B.2, we discuss the consistency of the non-constant trade elasticities with our theoretical framework and provide an extension of our gravity model.

4 Endogeneity

This section takes the identification of the causal effects of military alliances on trade a step further. We propose two approaches. First, we develop an instrumental variable based on common outside alliances and the Conley et al. (2012) test. Then, we perform dynamic propensity score matching and estimate the Differenced Average Treatment on the Treated (DATT) based on Couch and Placzek (2010).

4.1 Instrumental variable approach

We define our instrumental variable as the sum of common outside alliances. The intuition behind this is the domino-like spread of international agreements, as tested for RTAs or regionalism by Baldwin and Jaimovich (2012). Common allies present similar military interests and the inability of one country in the pair to use its alliances in a conflict against the other. Therefore, if countries i and j have many allies in common, it is highly probable they will develop an alliance together. Yet, we want to avoid capturing the inverse relation: an increase of the common outside alliances because of the signature of an alliance between countries i and j. For this purpose, we ignore variations in common outside alliances while country i and j are allied.[32] Thus, we can write the instrumental variable as follows:

$$IV_{ijt} = \begin{cases} \sum\limits_{k \neq i,j;t_{all}} (ALL_{ikt_{all}} * ALL_{kjt_{all}}), \text{ if } \exists t_{all}, t > t_{all} \ \& \ ALL_{ijt} = 1 \\ \\ \sum\limits_{k \neq i,j;t} (ALL_{ikt} * ALL_{kjt}), \text{ otherwise} \end{cases} \tag{6}$$

We exclude from the sum the alliance between country-pair ij – the country-pair of interest. The sum of common outside alliances is time-country-pair varying.

[32] Formally, as long as country i and j are allies, we set the value of the sum of their common outside alliances at the year of signature t_{all}

We use an OLS/PPML two-stage approach to keep the IV strategy comparable with the baseline and prevent log-linearisation under heteroscedasticity from biasing our estimates (Silva and Tenreyro, 2006). In the first stage, we use OLS to estimate the effect of the IV on the probability of signing an alliance. The predicted probability is used to compute the instrumented alliances. Then, we estimate with a PPML the effect of the instrumented alliances on bilateral exports. In the second stage, clustered standard-errors are bootstrapped.[33] Results are reported in table 4. The (instrumented) military alliance coefficient is strongly positive and significant.[34] Therefore, the IV estimation confirms the causal interpretation of the effects of military alliances on trade.[35]

Using a non-linear second-stage estimator may induce consistency issues. Lin and Wooldridge (2019) recommend a control-function approach to address this. Instead of using alliances' predicted probabilities, we include first-stage residuals as control. Results, reported in appendix table 16, confirm the OLS/PPML two-stage conclusions. In addition, the control-function approach allows testing the baseline's sensitivity to omitted variable concerns. The coefficient of the first-stage residuals is directly interpretable as the omitted variable bias addressed by the IV. Interestingly, focusing on defence pacts (column 2), the residuals' coefficient is significantly smaller, suggesting that defense pacts are less concerned by omitted variable issues than weak alliances. [36]

To respect the exclusion restriction, the signature of a military alliance with a third country k must not affect trade between i and j. We create two dummies to qualitatively test this assumption: $ALL^{out}_{i,-j,t}$ and $ALL^{out}_{-i,jt}$. $ALL^{out}_{i,-j,t}$ takes value 1 if the exporter has signed an alliance with any country other than j (i.e. an outsider) and zero otherwise. Similarly, $ALL^{out}_{-i,jt}$ takes value 1 if the importer has signed an alliance with any country other than i. Both $ALL^{out}_{i,-j,t}$ and $ALL^{out}_{-i,jt}$ can be estimated in the presence of exporter-year, importer-year and country-pair fixed effects since they are country-pair-year specific (Dai et al., 2014). Both dummies are introduced in equation (4) estimated by a PPML. We report the results in appendix table 17.[37] $ALL^{out}_{i,-j,t}$'s and $ALL^{out}_{-i,jt}$'s coefficients are not significantly different from zero. So, signing an alliance with an outsider does not affect exports from country i to j. This result points to the validity of the instrument's exogeneity assumption. However, the next section presents a plausible exogeneity test designed to support the validity of our results even in the

[33] Bootstrapping not only the second stage but the whole process does not affect results.

[34] We observe a higher coefficient compared with the baseline results. This is due to alliance selection induced by the IV. By targeting alliances included in an international network, we mechanically select defence pacts, which tends to increase coefficient (see section 3.3). Reduction in second-stage standard errors results simply from bootstrapping.

[35] Supplementary sensitivity tests are performed on the IV results. Results are robust to the inclusion of an interaction variable between distance (in log) and year, which controls for the globalisation dynamic. To avoid potential biases due to RTA endogeneity, we also run an estimation in which we instrument them by common outside RTAs ($\sum_{k \neq i,j;t} RTA_{ikt} * RTA_{kjt}$). This does not affect the instrumented alliance coefficient. We provide a 2SLS estimation in appendix table 18. The alliance coefficient is still positive and significant. Yet, the OLS and 2SLS gravity estimations induce biased coefficients and standard errors – due to log-linearisation under heteroscedasticity (Silva and Tenreyro, 2006) – and observation weights different to PPML (Sotelo, 2019).

[36] Furthermore, coefficients of instrumented alliances and defence pacts are very similar, which supports that defence pacts drive the estimated causal effect on trade.

[37] We run the same estimation in column two, but decompose the ALL_{ijt} dummies into defence pacts and weak alliances. This does not affect the results.

presence of a reasonable deviation from perfect alignment with the exclusion restriction assumption.

Table 4: Alliances and bilateral exports, IV

Dependent variable: exports	
Second stage	(1)
Estimator:	PPML
Instrument variable:	Common out. alliances
Alliance	0.655^a
	(0.026)
RTA	0.159^a
	(0.011)
Exporter x Year FE	yes
Importer x Year FE	yes
Dyadic FE	yes
No. observ.	320.666
First-stage	
Estimator:	OLS
Instrumented variable:	Alliance
Common out. alliances	0.056^a
	(0.007)
RTA	0.011
	(0.008)
Exporter x Year FE	yes
Importer x Year FE	yes
Dyadic FE	yes
No. observ.	320.712
KPW F-stat	67
KPW LM-stat	11

Note: OLS, Ordinary Least Squares; FE, Fixed effects. Dependent variable is exports from country i to country j at time t in millions of current dollars. Standard errors clustered at the exporter and importer levels are in parentheses. Second-stage standard errors are bootstrapped. a, b and c denote significantly different from 0 at the 1%, 5% and 10% level, respectively.

4.1.1 IV validity

The IV approach is based on the assumption of non-violation of the exclusion restriction. The common outside alliances can impact country i's exports to j only through enforcement of an alliance. We argue above that we have good reason to consider the restriction valid. Yet, we perform the *plausible exogeneity test* proposed by Conley et al. (2012)to properly address this point. The approach relaxes the assumption by allowing for ν, the correlation between the instrumental variable (common outside alliances) and errors, to deviate from zero. We then test whether the estimate of the instrumented variable (ALL_{ijt}) is robust to a variety of deviations from the exclusion restriction.

The union of confidence interval method calls first for the setting of the minimum (or maximum) value ν can take. To approximate this, we regress bilateral exports on the endogenous variable (ALL_{ijt})

and the instrumental variable (common outside alliances) with our standard set of fixed effects and controls. The coefficient associated with the IV represents an approximation of the degree of deviation from the exclusion restriction (Kippersluis and Rietveld, 2018). We obtain a small coefficient (-0.030). We then plug this degree of deviation into the plausible exogeneity test. In this way, we obtain the estimation of the (instrumented) alliance coefficient's upper and lower bounds under the relaxation of the exclusion restriction (see table 5).[38] The interval [0.388 ; 1.089] does not contain zero. Thus, we can safely argue that military alliances have an unambiguous positive causal effect on bilateral exports.

Table 5: Plausible exogeneity test

Dep var: exports	*Union of Confidence Interval estimations*		
Instrumented var.	ν -0.030	Min 95% CI	Max 95% CI
Alliance		0.408	1.106

Note: UCI based on the IV's ν coefficient from a regression of exports on interest variables and the IV

4.2 Differenced average treatment on the treated

In this section, we estimate the Differenced Average Treatment on the Treated (DATT) based on Couch and Placzek (2010). In the light of the results in section 3.3, we choose to increase the precision of the analysis by focusing on defence pacts. Therefore, we decompose variable ALL_{ijt} into *defence pacts* and *weak alliances*. Then, we estimate the defence pact DATT, using weak alliances as a control variable.

First, we perform dynamic propensity score matching. We define the propensity score $p(x_{ijt})_t$ as the likelihood of signing a defence pact conditional on a set of standard observable gravity variables. The propensity score is estimated for each separate year from 1967 to 2012. In this way, depending on the year, the variables are allowed to affect $p(x_{ijt})_t$ differently. Each treated country-pair observation is matched on $p(x_{ijt})_t$ with an untreated country-pair. A country-pair with no defence pact, but having had one in the past is never used as a control observation. We obtain a set of country-pair-year observations matched on $p(x_{ijt})_t$.[39]

[38] The Conley et al. (2012) estimator is based on the 2SLS estimator. We therefore have to log-linearise the gravity equation. We use the inverse hyperbolic sine transformation of exports to capture zero values. As discussed previously, log-linearisation under heteroscedasticity could bias the alliance coefficient. Nevertheless, estimated bounds are consistent with the OLS/PPML two-stage estimation which addresses this bias.

[39] See the appendix tables 19 and 20 for more details on dynamic propensity score matching. We evaluate the extent to which the matched treated and control groups are similar. We consider the standardised difference in means (B) and the variance ratio (R). To conclude that the groups are comparable, B must be inferior to 0.25, and R between 0.5 and 2 (Rubin, 2001; Stuart and Rubin, 2008). After matching, we obtain $B = 24.7\% < 25\%$ and $0.5 < R = 1.7 < 2$. Thus, in the set of matched observations, the control and treated groups are well balanced.

Second, in keeping with Couch and Placzek (2010), we estimate the equation:

$$X_{ijt} = exp(\sum_{k>k'} \delta^k D_{ijt}^k + \beta_2 Weak.ALL_{ijt} + \beta_3 RTA_{ijt} + \lambda_{it} + \lambda_{jt} + \lambda_{ij}) * \epsilon_{ijt} \qquad (7)$$

k' must be a year or period prior to the year of signature of a defence pact between countries i and j. D_{ijt}^k is a dummy variable equal to 1 if t is the k^{th} year after (or the k^{th} year before if -k) the signature of a defence pact between countries i and j. $Weak.ALL_{ijt}$ and RTA_{ijt} are dummy variables controlling respectively for the presence of a weak alliance and an RTA between countries i and j at time t. λ_{it}, λ_{jt} and λ_{ij} are respectively exporter-year, importer-year and exporter-importer fixed-effects. We use a PPML estimator.

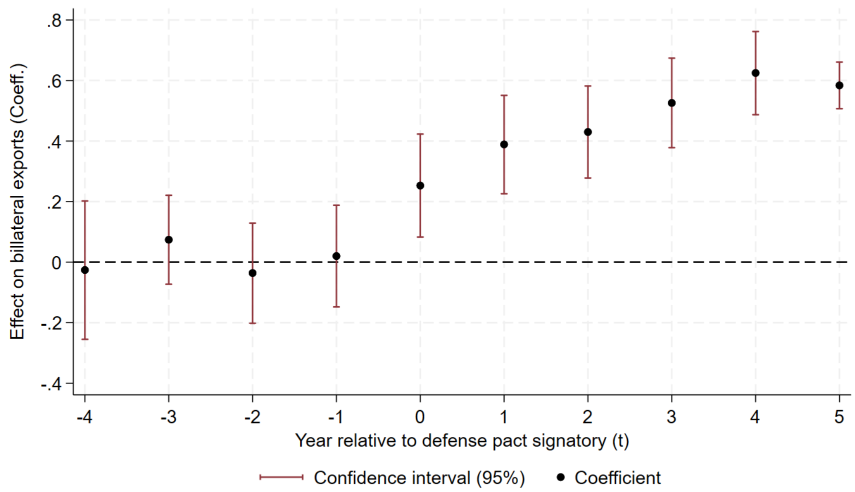
The DATT compares the difference in bilateral exports between years k and k' for a country-pair signing a defence pact during our period, indicated by $D_{ij} = 1$, to the difference in bilateral exports between years k and k' for a country-pair with no defence pact, indicated by $D_{ij} = 0$. The expected difference between the year-to-year difference in exports is estimated for the set of treated country-pairs relative to the matched set of non-treated pairs (Couch and Placzek, 2010).

The results of the DATT estimation are graphically represented in figure 6. More details are provided in appendix table 21. t_{def} is the date of the defence pact's signature. We choose $k' <= t_{def} - 5$. So, all k are estimated in comparison to $k' <= t_{def} - 5$. As soon as the treaty is signed, the defence pact has a positive and significant effect on exports from i to j. In the following years, the effect keeps growing to stabilize in $t_{def} + 5$. Given that t_{def} the date of signature and not of enforcement, this period can be interpreted as the time of adjustment required for the defence pact to become effective and fully operational. For $k >= t_{def} + 5$, the average estimated effect confirms our previous results: following their signature, defence pacts increase bilateral trade by 79%.[40] Moreover, for any $k < t_{def}$, we obtain a weak and insignificant coefficient. Therefore, the measured effect for any $k >= t_{def}$ is independent of any pre-trend.[41] Thus, in the light of the IV and DATT results, we can safely conclude that the defence pacts have an unambiguous gradual, positive and causal effect on trade.

[40] Since the data set is staggered, the DATT estimates only the entry effect (switches from 0 to 1). In previous estimation strategies, both entry and exit are considered.

[41] DATT results are also robust to the cold war. See the full discussion on the cold war in section D.

Figure 6: Dynamic effect of defence pacts on bilateral exports



5 From alliances to trade: the mechanism

This section analyses the mechanism behind the positive effect of alliances on trade.[42] We test the effect of alliances directly on insecurity and then on exports. Using data on conflict events, we define bilateral insecurity as:

$$INS_{ijt} = \sum(conflict\_event_{c_{ij},it}|c_{ij} = 0) * \sum(conflict\_event_{c_{ij},jt}|c_{ij} = 0) \quad (8)$$

the interacted country-time sums of conflict events excluding events with military cooperation between i and j.[43] This measure has evident advantages over the proxies presented in section 2.2.1. First, it presents a good country-pair-time variation. Second, it directly targets insecurity. Third, it captures the extensive margin of bilateral insecurity (if country i and j are exposed to insecurity) and the intensive margin (to what extent the country-pair is exposed).

We use two-stage estimations to look at the effect of defence pacts on exports by means of the variation of INS_{ijt} they enforce. For interpretation reasons, we use the inverse hyperbolic sine transformation of INS_{ijt}.[44] The set of fixed effects is the same as in the baseline.[45] Due to data limitations with respect to military events, we use a sub-sample starting in 1989. The results are displayed in table

[42]Most previous empirical papers focusing on the link between military alliances and trade consider that alliances affect exports because signatories are more inclined to reduce tariffs or sign RTAs with their allies (Long, 2003; Long and Leeds, 2006; Mansfield and Bronson, 1997). Yet, our baseline estimation includes RTAs as a control variable, while our analysis shows that our baseline results are barely sensitive to the inclusion of tariffs and RTA depth (see section 3). Therefore, this mechanism is excluded.

[43]Military cooperation events capture a reduction in insecurity rather than an increase. They therefore need to be excluded for an accurate measurement of bilateral insecurity.

[44]We test different functional forms. This does not affect the robustness of our results. In addition, given that on average $INS_{ijt} = 14622$, we are not exposed to approximation errors using the inverse hyperbolic sine transformation (Bellemare and Wichman, 2020).

[45]Note that this implies that the independent level of insecurity of country j or in country j at time t are controlled and cannot bias our estimates.

6.[46] Column 1 presents the defence pact dummy variable. To further address possible endogeneity, we replace it in column 2 with common outside defence pacts.[47]

In the first-stage estimations, we observe a strong negative effect of defence pacts on bilateral insecurity. In column 1, we estimate that the enforcement of a defence pact reduces bilateral insecurity by 69%. Then, in the second-stage results, we observe in both specifications that the reduction in bilateral insecurity enforced by defence pacts (or common outside defence pacts) has a significant positive effect.[48] Thus, we estimate that, by reducing bilateral insecurity, defence pacts raise exports by 49.5%. In this sub-sample, this is equivalent to almost 100% of the total defence pact effect. Therefore, these results strongly support both the validity and prevalence of the insecurity mechanism.

Alternatively, we test the insecurity mechanism estimating the effect of defence pacts on exports through the variation of $\sum coop.mil.ev._{ijt}$, the sum of the country-pair's cooperative military events. Contrary to INS_{ijt}, this does not directly target insecurity, but the enforcement of the military cooperation that reduces insecurity. Yet, we identify a purely bilateral variation. We report the two-stage estimations in appendix table 22. Results confirm our conclusions.[49]

[46] As in section 4.1, we use a two-stage OLS/PPML estimator to prevent log-linearisation under heteroscedasticity from biasing our results (Lin and Wooldridge, 2019; Silva and Tenreyro, 2006).

[47] We construct this variable in the same way as the IV in section 4.1 but restricting it to defence pacts.

[48] The bilateral insecurity coefficient is negative since it expresses the effect of a *rise* in insecurity.

[49] $\sum coop.mil.ev._{ijt}$ measures the opposite of INS_{ijt}. Therefore, the second-stage coefficient is positive: defence pacts make for a sharp increase in military cooperation, which makes for a sharp increase in trade.

Table 6: The bilateral insecurity reduction

Second stage	(1)	(2)
Estimator:	PPML	
Dependent variable:	Exports	
Bilateral insecurity	-0.346^{a}	-0.311^{a}
	(0.049)	(0.045)
Weak alliance	-0.227	-0.286^{c}
	(0.152)	(0.149)
RTA	0.040	0.041
	(0.036)	(0.033)
Exporter x Year FE	yes	yes
Importer x Year FE	yes	yes
Dyadic FE	yes	yes
No. observ.	167.304	167.304
First-stage		
Estimator:	OLS	
Dependent variable:	Bilateral insecurity	
Defence pact	-1.162^{a}	
	(0.148)	
Common out. def. pacts		-0.076^{a}
		(0.009)
Weak alliances	-0.213	-0.130
	(0.243)	(0.242)
RTA	-0.165^{a}	-0.168^{a}
	(0.060)	(0.060)
Exporter x Year FE	yes	yes
Importer x Year FE	yes	yes
Dyadic FE	yes	yes
No. observ.	167.328	167.328

Note: OLS, Ordinary Least Square; PPML, Poisson Pseudo Maximum Likelihood; FE, Fixed effects. The panel starts in 1989. Common defence pacts sum all external partners for which country i and j both have a defence pact. Robust standard errors clustered at country-pair level are in parentheses. Second-stage standard errors are bootstrapped. a, b and c denote significantly different from 0 at the 1%, 5% and 10% level, respectively

6 General equilibrium and welfare implications

Following the theoretical model presented in section 1 and detailed in section B, we can now analyse the general equilibrium ramifications of signing a military alliance. Four equations characterise our model:

$$X_{ij} = N_i \overline{\alpha}_i^{-\theta} w_i^{-\theta} \frac{X_j}{\Phi_j^{-\theta}} \tau_{s,ij}^{-\theta} f_{s,ij}^{-[\frac{\theta}{\sigma-1}-1]} T_{n\neq s,ij}^{-\theta} F_{n\neq s,ij}^{-[\frac{\theta}{\sigma-1}-1]} \tag{9}$$

$$\Pi_i^{-\theta} = \sum_j \frac{\tau_{s,ij}^{-\theta} f_{s,ij}^{-[\frac{\theta}{\sigma-1}-1]} T_{n\neq s,ij}^{-\theta} F_{n\neq s,ij}^{-[\frac{\theta}{\sigma-1}-1]}}{\Phi_j^{-\theta}} X_j \tag{10}$$

$$\Phi_j^{-\theta} = \sum_i \frac{\tau_{s,ij}^{-\theta} f_{s,ij}^{-[\frac{\theta}{\sigma-1}-1]} T_{n\neq s,ij}^{-\theta} F_{n\neq s,ij}^{-[\frac{\theta}{\sigma-1}-1]}}{\Pi_i^{-\theta}} Y_i \tag{11}$$

$$\overline{\alpha}_i w_i = (\frac{Y_i}{\Pi_i^{-\theta} N_i})^{-\frac{1}{\theta}} \tag{12}$$

Equation 9 is the structural gravity relation presented in section 1 above. Π_i and Φ_j are respectively the outward and inward multilateral resistance terms. [50] Y_i is country i's total output and $\overline{\alpha}_i w_i$ the maximum factory gate price – the marginal cost at the lowest productivity to which a firm can drop in country i.

In this frame, in keeping with Arkolakis et al. (2012), we can summarise the welfare effect of a change in insecurity costs due to military alliance enforcement in a few parameters. Welfare is defined as the real revenue ($\frac{Y}{\Phi}$). Thus, any change in welfare follows the equality:

$$dln(W_j) = dln(Y_j) - dln(\Phi_j) \tag{13}$$

Military alliances contend with bilateral and reciprocal shocks on variable and fixed trade costs sensitive to insecurity. Based on such shocks, we can desegregate country j changes in welfare as follows:

$$dln(W_j) = dln(w_j) - \sum_{i=1}^{n} \psi_{ij} \left[dln(w_i) + dln(\tau_{s,ij}) + (\frac{1}{\sigma-1} - \frac{1}{\theta}) dln(f_{s,ij}) \right] \tag{14}$$

Therefore, welfare changes induced by military alliances can be derived from a system of a few parameters: initial trade shares (ψ_{ij}), wages (w_j and w_i), insecurity costs ($\tau_{s,ij}$ and $f_{s,ij}$)[51] and

[50] Φ_j in eq. 11 is obtained by replacing in eq. 31 $N_i \overline{\alpha}_i^{-\theta} w_i^{-\theta} = \frac{Y_i}{\Pi_i^{-\theta}}$, which is derived from the market clearance.

[51] More precisely, since we are looking at a change in insecurity costs produced by military alliances, we have $dln(\tau_{ij}(s)) = \varepsilon_{ln(\tau_{s,ij})}(ALL_{ijt})$ and $dln(f_{ij}(s)) = \varepsilon_{ln(f_{s,ij})}(ALL_{ijt})$, respectively the variable and fixed insecurity costs alliance elasticities. Thus, for the empirical application in the next section, we do not need to identify changes in insecurity costs, but only changes in alliances.

elasticities of trade and substitution (θ and σ).

6.1 Full endowment GE: empirical application

This welfare system can be solved using the Anderson and Yotov (2016) methodology and its extension in Anderson et al. (2018). We first estimate the partial equilibrium (eq. 9)[52] to capture military alliance elasticity and estimates of bilateral trade costs.[53] We develop our analysis for the year 2012 – the last year in our panel. To retrieve baseline multilateral resistance, we estimate the gravity equation for 2012 imposing bilateral trade costs and elasticities from the previous step.[54] We define our counterfactual as the absence of alliance. To obtain the counterfactual multilateral resistance terms, we again estimate the constrained gravity equation setting $ALL_{ijt} = 0$. We then determine the endogenous change in output and expenditures: $\widetilde{X_i^c} = \frac{\tilde{w_i^c}}{\tilde{w_i}} X_i$ and $\widetilde{Y_i^c} = \frac{\tilde{w_i^c}}{\tilde{w_i}} Y_i$. The Change in wage $(\frac{\tilde{w_i^c}}{\tilde{w_i}})$ is captured directly by changes in maximum factory gate prices (eq. 12).[55] The calibration of trade elasticity θ plays a crucial role in the estimation of both prices and welfare. Including tariffs in the standard structural gravity estimation, we directly estimate $\theta = 3.7$.[56] Finally, we can quantify the General Equilibrium effect of military alliances enforced in 2012 as the percentage difference between the baseline and the counterfactual scenarios.[57]

6.2 Results

Results per country are reported in table 7.[58] Military alliances improve signatories' exports and welfare, while non-members experience losses. The biggest winners in terms of welfare also see sharp increases in their exports, but not the largest ones. These countries are small, developed countries with military alliances with many partners, such as Belgium (16.28%), the Netherlands (+16.20%)

[52]Given that this is a general equilibrium case, intranational trade flows are included. See section 3.2 for details on this specification and comparison with our baseline.

[53]Time-varying trade costs are derived from controls, while time-invariant trade costs are captured by country-pair fixed effects. Some exporter-importer fixed effects are dropped due to convergence issues. These missing effects are replaced by regressing for 2012 the estimates of exporter-importer fixed effects on gravity variables and country fixed effects.

[54]Using a PPML estimator, we can directly recover empirical expressions of the multilateral resistance terms (Fally, 2015). Yet to solve the system, we need to normalise one of the multilateral resistances. We choose to normalise Germany's importer multilateral resistance term so that $\widetilde{\Phi_0} = 1$. Therefore we can derive country i's and j's multilateral resistance from $\widetilde{\Pi_i} = (\frac{Y_i X_0}{exp(\widetilde{\lambda_i})})^{-\frac{1}{\theta}}$ and $\widetilde{\Phi_j} = (\frac{X_j}{X_0 exp(\widetilde{\lambda_j})})^{-\frac{1}{\theta}}$.

[55]Given that $\overline{\alpha}$ and N_i are fixed, we have $\frac{\tilde{w_i^c}}{\tilde{w_i}} = (\frac{exp(\tilde{\lambda_i})^c X_0}{exp(\tilde{\lambda_i}) X_0^c})^{-\frac{1}{\theta}}$. Variations in expenditures and outputs trigger new changes in multilateral resistance, which impacts outputs and expenditures and so forth. Translating these variations into changes in exports, we iterate the estimation process until maximum prices converge.

[56]The results of the trade elasticity estimation are reported in table 23. For robustness reasons, we also perform the general equilibrium analysis using two alternative calibrations of θ. In keeping with Head et al. (2014) and Melitz and Redding (2013) we calibrated $\theta = 4.25$, and in keeping with Anderson et al. (2018) $\theta = 6$. Mechanically, the higher θ, the smaller the welfare variations. Nonetheless, our conclusions are robust to these alternative calibrations of θ.

[57]Changes in exports and welfare are simply: $\frac{X_j}{X_j^c} = \frac{\sum_j X_{ij} - \sum_j \widetilde{X_{ij}^c}}{\sum_j \widetilde{X_{ij}^c}}$ and $\frac{W_j}{W_j^c} = \frac{Y_j/\widetilde{\Phi_j} - \widetilde{Y_j^c}/\widetilde{\Phi_j^c}}{\widetilde{Y_j^c}/\widetilde{\Phi_j^c}}$).

[58]Minor additional sensitivity checks are performed. We replicate the analysis while controlling for globalisation as in table 12 column 8. We also differentiate between defence pacts and weak alliances, removing only defence pacts. In both checks, results are barely affected.

and Iceland (+13.12%). We then find a mix of developed and middle-income countries also intensively involved in military alliance treaties, such as Hungary (+11.03%), Mexico (+10.71%), Venezuela (+9.75%), Denmark (+8.75%), Canada (+8.74%) and Argentina (+6.06%). Military alliances are less impactful for large developed economies, but still bring non-negligible welfare gains: France (+5.34%), Great Britain (+5.21%), Germany (+4.59%) and the USA (+2.42%). On the other hand, the biggest losers are small countries signatories to few military alliance treaties whose close partners have signed multiple treaties. This is characteristic of a general equilibrium trade diversion mechanism. Here, we can cite Ireland (-3.30%), Nigeria (-2.85%), Malta (-2.44%), Cyprus (-2.35%) and Sweden (-2.21%).[59] In addition, our results are robust to the inclusion of heterogeneous elasticities.[60]

6.3 Scenarios analysis

Interstate tensions have been exacerbated in recent years. The war in Ukraine has brought conflict back to Europe with new threats and strategic interests that are shaking the post-Cold War balance. The future may bring a substantial reshaping of military alliances, especially among European countries. Using our general equilibrium approach, this section analyses the ramifications on countries' welfare of potential alliance network disruption.[61] We define three scenarios: i) the expansion of NATO to neutral European countries and partner nations, ii) NATO breaking with Eastern countries, and iii) the creation of a new Eastern bloc.[62]

Results per country are presented in the appendix table 25. Any scenario's GE effect is insufficient to impact large or unconcerned economies. For countries like the US, France, UK and Germany, maintaining NATO in the East or expanding the treaty to their closest partners has noticeable, but not major economic outcomes. On the other hand, for countries targeted as potential switchers, the choice to switch would have drastic repercussions. Leaving NATO would severely reduce their real revenue, which a new Eastern bloc could temper but not offset. Moreover, for still-neutral countries

[59] Note that losers may experience (very small) gains in exports. This is the result of an increase in their inward and outward multilateral resistance terms associated with a drop in their factory gate prices. The increase in expenditure by winners causes non-signatory countries to see trade diverted from their internal market to the winners. This diversion offsets (or overcomes) the decrease in their exports due to the fall in output.

[60] We showed in section 3.5 5 that, given that trade elasticities are not constant, alliances have a very heterogeneous effect. We adjust our GE analysis to include non-constant trade elasticities. Basically, we allow trade elasticities to be country-pair-year specific and the welfare elasticity to be country-year specific. Theoretically, the above-presented GE structure and welfare system still hold. After estimating the average effect on military alliances, we introduce the distortion in coefficients derived from the quantile estimation in section 3.5. Then, we determine the country-specific welfare elasticities. We calibrate the average elasticity at 3.7 and apply the country-year specific distortion. In keeping with our model extension in section B.2, a good proxy can be directly derived from the distribution of the country-year averages of the alliance elasticities. We present the new GE results in appendix table 24. On the whole, the introduction of non-constant elasticities tempers the welfare variations. Nonetheless, our conclusions are barely affected.

[61] We do not observe Russia or Ukraine's intranational trade, which excludes them from the analysis. Therefore, we do not aim to study the impact of war on the two belligerents directly, but the ramifications of new military relationships between countries indirectly involved.

[62] We assume that i) Sweden, Finland, Japan, Austria, Australia, Ireland, South Korea, New Zealand, Colombia, Malta, Switzerland, Pakistan and Serbia-Montenegro become full NATO members, ii) Albania, Bulgaria, China, Finland, Sweden, Hungary, Pakistan, Poland, Romania, Serbia-Montenegro and Turkey leave NATO or terminate any military alliance with its members, and iii) that after leaving NATO, they form one common military alliance together.

– especially in Europe – joining NATO (i.e. the closest and largest alliance network) would bring remarkable welfare gains. For example, Bulgaria would increase its real revenue by 10.27%, Finland by 11.51%, Sweden by 15.73%, Switzerland by 17.13% and Ireland by 24.58%.

7 Conclusion

This paper provides a systematic analysis of the impact of military alliances on trade using a panel of 6,972 country pairs from 1967 to 2012. Taking a structural gravity approach, we show that military alliances have a strong positive causal effect on bilateral exports. Namely, our baseline specification shows that the enforcement of a military alliance engenders an increase of 60%. We perform numerous sensitivity tests and show that results are robust to a variety of alternative specifications. However, this average effect is highly heterogeneous across country pairs, depending on their exposure to trade margins, the nature of the treaties and the non-constancy of trade elasticities. Furthermore, an instrumental variable approach and a DATT approach confirm the causal interpretation of our results. We confirm the validity and prevalence of the insecurity mechanism: military alliances increase trade by reducing bilateral insecurity. Finally, we perform a general equilibrium analysis to quantify the welfare effect of alliances. Building a counterfactual for 2012, we show that intensive involvement in the signature of military alliances brings significant welfare gains while being neutral induces marked losses. We then analyse different scenarios to demonstrate that reshaping the military alliance network in response to the war in Ukraine could have considerable welfare ramifications on the economies concerned.

Our findings have important scientific and policy implications. First, they point up the need to consider the specific role of security and international military relations to understand trade and globalisation. Second, they show the efficiency of military alliances, particularly defence pacts, at guaranteeing the safety and inter-state cooperation required for economic agents to trade. The unambiguous welfare gains that alliances bring their members should give policymakers strong incentives to promote the signature of such treaties. Although some may fear that they create relations of domination between nations, our findings suggest on the contrary that they bring more favourable welfare gains to small economies.

Table 7: GE Exports and Welfare

Country (Iso3)	Exports	Real revenues	Country (Iso3)	Exports	Real revenues
ALB	2.34	-0.38	ISL	29.96	13.12
ARG	21.54	6.06	ISR	-2.66	-1.84
AUS	4.16	0.46	ITA	35.10	3.53
AUT	-1.62	-1.93	JPN	7.37	0.40
BEL	19.95	16.28	KEN	-5.54	-0.34
BGD	-0.36	-0.99	KOR	5.45	0.61
BGR	9.02	1.68	LKA	-2.06	-0.91
BOL	29.58	11.60	MAR	-0.15	-0.52
BRA	20.02	0.74	MEX	28.92	10.71
CAN	44.41	8.74	MLT	1.02	-2.44
CHE	-1.05	-1.89	MYS	0.63	-0.71
CHL	11.76	6.11	NGA	1.44	-2.85
CHN	-1.14	-0.13	NLD	17.80	16.20
CIV	13.14	0.30	NOR	20.31	6.61
CMR	0.20	-1.25	NZL	-0.51	-0.44
COL	38.52	3.32	PAK	25.09	5.19
CYP	-1.04	-2.35	PER	25.07	2.79
DEU	26.66	4.59	PHL	4.21	1.70
DNK	26.85	8.75	POL	61.18	5.11
DZA	-0.45	-1.02	PRT	56.90	6.94
ECU	31.62	6.26	ROM	2.24	-0.19
EGY	-3.64	-0.68	SAU	0.38	-0.96
ESP	45.81	4.14	SER	-5.52	-0.55
FIN	-2.55	-1.28	SGP	-0.39	-0.38
FRA	45.15	5.34	SWE	-2.32	-2.21
GAB	1.38	3.66	THA	0.31	-0.64
GBR	41.19	5.21	TUN	0.22	-0.77
GRC	51.44	4.98	TUR	60.8	2.53
HKG	0.02	-1.36	URY	37.82	3.86
HUN	26.82	11.03	USA	65.90	2.42
IDN	0.14	-0.32	VEN	18.17	9.75
IND	-2.36	-0.20	VNM	0.05	-0.45
IRL	1.25	-3.30			

Note: The real revenue is our measure of welfare. All numbers are percentage variations.

References

Anderson, J. E., Larch, M., and Yotov, Y. V. (2018). GEPPML: General equilibrium analysis with PPML. *The World Economy*, 41(10):2750–2782. Number: 10.

Anderson, J. E. and Marcouiller, D. (2002). Insecurity and the pattern of trade: An empirical investigation. *Review of Economics and statistics*, 84(2):342–352.

Anderson, J. E. and Van Wincoop, E. (2004). Trade costs. *Journal of Economic literature*, 42(3):691–751.

Anderson, J. E. and Yotov, Y. V. (2016). Terms of trade and global efficiency effects of free trade agreements, 1990–2002. *Journal of International Economics*, 99:279–298.

Arkolakis, C., Costinot, A., and Rodríguez-Clare, A. (2012). New trade models, same old gains? *American Economic Review*, 102(1):94–130.

Baier, S. L. and Bergstrand, J. H. (2007). Do free trade agreements actually increase members' international trade? *Journal of International Economics*, 71(1):72–95. Number: 1.

Baier, S. L., Yotov, Y. V., and Zylkin, T. (2019). On the widely differing effects of free trade agreements: Lessons from twenty years of trade integration. *Journal of International Economics*, 116:206–226.

Baldwin, R. and Jaimovich, D. (2012). Are free trade agreements contagious? *Journal of international Economics*, 88(1):1–16.

Behrens, K., Ertur, C., and Koch, W. (2012). 'dual'gravity: Using spatial econometrics to control for multilateral resistance. *Journal of Applied Econometrics*, 27(5):773–794.

Bellemare, M. F. and Wichman, C. J. (2020). Elasticities and the inverse hyperbolic sine transformation. *Oxford Bulletin of Economics and Statistics*, 82(1):50–61.

Blomberg, S. B. and Hess, G. D. (2006). How much does violence tax trade? *The Review of Economics and Statistics*, 88(4):599–612.

Carrère, C., Mrázová, M., and Neary, J. P. (2020). Gravity Without Apology: the Science of Elasticities, Distance and Trade. *The Economic Journal*, 130(628):880–910. Number: 628.

Chaney, T. (2008). Distorted Gravity: The Intensive and Extensive Margins of International Trade. *American Economic Review*, 98(4):1707–1721. Number: 4.

Conley, T. G., Hansen, C. B., and Rossi, P. E. (2012). Plausibly exogenous. *Review of Economics and Statistics*, 94(1):260–272.

Coppedge, M., Gerring, J., Knutsen, C. H., Lindberg, S. I., Teorell, J., Altman, D., Bernhard, M., Fish, M. S., Glynn, A., Hicken, A., et al. (2022). V-dem [country–year/country–date] dataset v12. varieties of democracy (v-dem) project. *V-Dem Institute, Department of Political Science, University of Gothenburg, Gothenburg.*

Couch, K. A. and Placzek, D. W. (2010). Earnings Losses of Displaced Workers Revisited. *American Economic Review*, 100(1):572–589. Number: 1.

Dai, M., Yotov, Y. V., and Zylkin, T. (2014). On the trade-diversion effects of free trade agreements. *Economics Letters*, 122(2):321–325.

De Chaisemartin, C. and d'Haultfoeuille, X. (2020). Two-way fixed effects estimators with heterogeneous treatment effects. *American Economic Review*, 110(9):2964–96.

De Chaisemartin, C. and D'Haultfoeuille, X. (2022). Difference-in-differences estimators of intertemporal treatment effects. *National Bureau of Economic Research*.

de Saint-Vaulry, A. (2008). *Base de données CHELEM-commerce international du CEPII*. CEPII.

De Sousa, J., Mayer, T., and Zignago, S. (2012). Market access in global and regional trade. *Regional Science and Urban Economics*, 42(6):1037–1052.

Fally, T. (2015). Structural gravity and fixed effects. *Journal of International Economics*, 97(1):76–85.

Feenstra, R. C. (2015). *Advanced international trade: theory and evidence*. Princeton university press.

Findlay, R. and O'Rourke, K. H. (2009). *Power and plenty: trade, war, and the world economy in the second millennium*, volume 30. Princeton University Press.

Fontagné, L. and Santoni, G. (2021). Gvcs and the endogenous geography of rtas. *European Economic Review*, 132:103656.

Gibler, D. M. (2008). *International military alliances, 1648-2008*. CQ Press.

Glick, R. and Taylor, A. M. (2010). Collateral damage: Trade disruption and the economic impact of war. *The Review of Economics and Statistics*, 92(1):102–127.

Head, K. and Mayer, T. (2014). Gravity Equations: Workhorse,Toolkit, and Cookbook. In *Handbook of International Economics*, volume 4, pages 131–195. Elsevier.

Head, K., Mayer, T., and Ries, J. (2010). The erosion of colonial trade linkages after independence. *Journal of International Economics*, 81(1):1–14. Number: 1.

Head, K., Mayer, T., and Thoenig, M. (2014). Welfare and trade without pareto. *American Economic Review*, 104(5):310–16.

Helpman, E., Melitz, M., and Rubinstein, Y. (2008). Estimating trade flows : Trading partners and trading volumes. *Quarterly Journal of Economics*, page 47.

Hofmann, C., Osnago, A., and Ruta, M. (2017). Horizontal depth: A new database on the content of preferential trade agreements. Policy Research Working Paper 7981, World Bank Group, Washington, DC.

Kippersluis, H. V. and Rietveld, C. A. (2018). Beyond plausibly exogenous. *Econometrics Journal*, 21(3):316–331.

Kitazawa, Y. (2012). Hyperbolic transformation and average elasticity in the framework of the fixed effects logit model. *Theoretical Economics Letters*.

Lin, W. and Wooldridge, J. M. (2019). Testing and correcting for endogeneity in nonlinear unobserved effects models. In *Panel Data Econometrics*, pages 21–43. Elsevier.

Long, A. G. (2003). Defense Pacts and International Trade. *Journal of Peace Research*, 40(5):537–552. Number: 5.

Long, A. G. and Leeds, B. A. (2006). Trading for security: Military alliances and economic agreements. *Journal of Peace Research*, 43(4):433–451.

Machado, J. A. and Silva, J. S. (2019). Quantiles via moments. *Journal of Econometrics*, 213(1):145–173.

Mansfield, E. D. and Bronson, R. (1997). Alliances, Preferential Trading Arrangements, and International Trade. *American Political Science Review*, 91(1):94–107. Number: 1.

Martin, P., Mayer, T., and Thoenig, M. (2008). Make trade not war? *The Review of Economic Studies*, 75(3):865–900.

Mayer, T., Santoni, G., and Vicard, V. (2023). The cepii trade and production database. Technical report.

Melitz, M. J. (2003). The Impact of Trade on Intra-Industry Reallocations and Aggregate Industry Productivity. *Econometrica*, 71(6):1695–1725. Number: 6.

Melitz, M. J. and Redding, S. J. (2013). Firm heterogeneity and aggregate welfare. *National Bureau of Economic Research Working Paper*, (No. 1891).

Mrázová, M. and Neary, J. P. (2017). Not so demanding: Demand structure and firm behavior. *American Economic Review*, 107(12):3835–74.

Pemstein, D., Marquardt, K. L., Tzelgov, E., Wang, Y.-t., Medzihorsky, J., Krusell, J., and Römer, J. v. (2022). The v-dem measurement model: Latent variable analysis for cross-national and cross-temporal expert-coded data (v-dem working paper no. 21). *University of Gothenburg: Varieties of Democracy Institute*.

Redding, S. and Venables, A. J. (2004). Economic geography and international inequality. *Journal of international Economics*, 62(1):53–82.

Rohner, D., Thoenig, M., and Zilibotti, F. (2013). War signals: A theory of trade, trust, and conflict. *Review of Economic Studies*, 80(3):1114–1147.

Rubin, D. B. (2001). Using propensity scores to help design observational studies: application to the tobacco litigation. *Health Services and Outcomes Research Methodology*, 2(3-4):169–188.

Sandkamp, A., Stamer, V., and Yang, S. (2022). Where has the rum gone? the impact of maritime piracy on trade and transport. *Review of World Economics*, pages 1–28.

Silva, J. and Kemp, G. C. (2016). Partial effects in fixed effects models. *London Stata Users Gr Meet.*

Silva, J. S. and Tenreyro, S. (2006). The log of gravity. *The Review of Economics and statistics*, 88(4):641–658.

Sotelo, S. (2019). Practical aspects of implementing the multinomial pml estimator. Technical report, Mimeo.

Stone, R., B. J. B. M. (1963). Input-output relationships, 1954-1966. In *A Programme for Growth*, volume 3. Chapman and Hall.

Stuart, E. A. and Rubin, D. B. (2008). Best Practices in Quasi–Experimental Designs: Matching Methods for Causal Inference. In *Best Practices in Quantitative Methods*, pages 155–176. SAGE Publications, Inc.

Sundberg, R. and Melander, E. (2013). Introducing the ucdp georeferenced event dataset. *Journal of Peace Research*, 50(4):523–532.

Tinbergen, J. (1962). Shaping the world economy; suggestions for an international economic policy.

Weidner, M. and Zylkin, T. (2021). Bias and consistency in three-way gravity models. *Journal of International Economics*, 132:103513.

Yu, S., Beugelsdijk, S., and de Haan, J. (2015). Trade, trust and the rule of law. *European Journal of Political Economy*, 37:102–115.

Appendices

A Military alliance treaties

A.1 Overview

Military alliances are international treaties designed to develop international military cooperation policies. As defined by Gibler (2008) military alliances can be divided into four categories depending on the degree of restriction and involvement of signatories. First, the *military entente* implies a diplomatic exchange of information among members before taking any military decision. Second, the *neutrality pact* specifies that signatories must stay neutral in the event of a conflict involving one party to the pact. Third, the *non-aggression pact* states that signatories cannot declare war or engage in military action against treaty members. Fourth, the *defence pact* is where signatories agree on collective but centralised military management. It does not deny members' sovereignty, but enforces strong military cooperation in areas that matter to the signatory countries. Therefore, defence pacts reach a highly specific level of cooperation. While the first three categories of alliance mainly describe different international policies to keep peace between members, the defence pacts imply military cooperation to protect signatories from outside threats.

A.2 Defence pacts examples

The most famous defence pact is the *North Atlantic Treaty* (NATO). Signed on the fourth of April 1949 and still in force, it concerns most North American and Western European countries and was later expanded to a number of Eastern European nations. Created to ensure protection against the USSR and its satellites, it implies strong and centralised military cooperation. It was also designed with important economic and institutional objectives in mind such as economic collaboration, free institutions and stable well-being.[63] Yet, to pursue these objectives, the treaty specifies only military cooperation policies. Moreover, the ninth article includes the creation of a central council in charge of compliance with the alliance's constraints, organisation and objectives.

[63] "Article 2. The Parties will contribute toward the further development of peaceful and friendly international relations by strengthening their free institutions, bringing about a better understanding of the principles upon which these institutions are founded, and promoting conditions of stability and well-being. They will seek to eliminate conflict in their international economic policies and will encourage economic collaboration between any or all of them."

The *Treaty on Collective Security* (TCS) was signed by former USSR states in 1992 and is still in force. It was signed in recognition of the inability of the Commonwealth of Independent States to provide the required economic and commercial prosperity among members despite the tariff liberalisation it includes. The TCS aims to achieve members' trade objectives by enabling lasting stability and security throughout the region due to common and centralised management of military matters. It also specifies the creation of a collective security council in charge of defence decisions, armed forces coordination and the application of the treaty's purposes.

The *Joint Defense and Economic Cooperation Treaty* between the States of the Arab League was created in 1951 by Egypt, Lebanon, Saudi Arabia, Syria, Yemen Arab Republic and Iraq, and joined later by a number of Arab states. Even though the treaty has encountered problems due to internal tensions over members' relations with Israel, it is still in force and was reinvigorated following the USA intervention in Iraq. Once again, it concerns close military cooperation and the creation of a centralised council. It also contains explicit economic and trade objectives and aims to favour the development and trade of signatory countries.[64] Moreover, to assist the first council, a second council in charge of economic issues was created. Yet, this council cannot propose or enforce standard trade liberalisation policies such as tariff reductions.

The *Defense Pact of the African and Malagasy Union* was signed in 1961 by twelve French-speaking sub-Saharan countries to protect themselves from both internal and external interference. Nonetheless, the pact did not last long since it was terminated in 1964. The core of the treaty was the enforcement of peace and stability in the region and military cooperation by members. Yet, it also introduced a mandatory contribution to the development of free institutions, well-being and economic collaboration. Similarly to the previous example, international security was also considered here as a necessity for economic development. Moreover, the treaty stipulated the creation of a central council to take decisions with a two-thirds majority of pact members regarding the alliance's procedures.

The last example of defence pact is the *Treaty Instituting the Arab-Maghreb Union*. It was signed in 1989 by North African and Arab countries and is still in effect today. The agreement covers common management of defence and stability matters, the creation of a presidential council to centralise decisions and a judicial body to ensure the legal enforcement of decisions. A striking point in our case is the explicit objective of trade liberalisation and free movement of persons, services, capital and goods.[65] The treaty's third article also includes a clear objective of common economic and industrial

[64] "Article 7. In order to fulfil the aims of this Treaty and to bring about security and prosperity in Arab countries and in an effort to raise the standard of life therein, the contracting States undertake to collaborate for the development of their economic conditions, the exploitation of their natural resources, the exchange of their respective agricultural and industrial products, and generally to organise and coordinate their economic activities and to conclude the necessary inter-Arab agreements to realise such aims."

[65] "Article Two. The Union aims at:(...) - Working gradually towards achieving free movement of persons and transfer of services, goods and capital among them."

development for members.

A.3 Weaker alliances examples

A first example is the neutrality pact between Chile and Argentina called the *Treaty of Peace and Friendship between Chile and Argentina*. It was driven by increasing border disputes between the two states since 1970 and negotiations to settle them. It was finally signed in 1984 and is still in force today. It lays down military and economic objectives, especially maritime goals, and the establishment of a commission. Yet, contrary to a defence pact, it does not imply close cooperation between states but merely a frame to avoid armed conflicts and agree on each state's sovereignty.

The *Treaty of Peace, Friendship and Cooperation between India and the Soviet Union*, coded as a non-aggression pact, was signed in 1971 and was in effect until 1991. The treaty provides for respect for members' sovereignty and borders and an absence of interference in any domestic affairs. Moreover, it expresses the importance of economic cooperation and trade. Yet, it provides for no supra-national institution to be created. The agreement provides for a guarantee from participants, but neither centralised military cooperation nor the ambitions of a defence pact.

The *treaty of Friendship and Co-operation between France and Russia*, an entente pact, was signed by Francois Mitterrand and Boris Yeltsin in 1992. It is still in force today without any objection having ever been raised by either state. The main goal of the agreement is to ensure immediate consultation between signatories in the event of security issues or important diplomatic decisions. It also stipulates institutional and economic objectives such as the development of the manufacturing sectors, the promotion of democratic institutions and the facilitation of the movement of capital, persons and goods. Yet, there is no further military cooperation, no guarantee of peace, and no supra-national institution to ensure the application of the treaty.

B Theoretical model

B.1 Standard model

We start with a constant elasticity of substitution utility function. Consumers in each country maximise their utility by consuming $q(\omega)$ units of each differentiated good ω, noted by the following function:

$$U = [\int_{\Omega} q(\omega)^{\frac{\sigma-1}{\sigma}} d\omega]^{\frac{\sigma}{\sigma-1}} \tag{15}$$

where q is the consumed quantity, and σ the elasticity of substitution between two varieties of goods. After maximisation under the revenue constraint, we can define the consumed quantity for a specific variety as:

$$q(\omega) = p(\omega)^{-\sigma} X(\int_{\Omega} p_{\omega''}^{1-\sigma})^{\frac{\sigma-1}{1-\sigma}} \tag{16}$$

and firm ω's revenue:

$$x(\omega) = X_j(\frac{p(\omega)}{P_j^{\frac{1}{1-\sigma}}})^{1-\sigma} \tag{17}$$

where $x(\omega) = q(\omega)p(\omega)$, X_j the total revenue of country j, and $P_j = \int_{\Omega} p_{\omega''}^{1-\sigma}$ the Dixit-Stiglitz price index. Thus, the total consumed value in country j of goods variety ω can be understood as the share of the country's total revenue allocated to the consumption of variety ω.

Next, we consider firm productivity level φ such that marginal cost $\alpha = 1/\varphi$. In keeping with Helpman et al. (2008), we assume that the distribution of firm productivity γ in each country follows a Pareto distribution G with:

$$G_i(\alpha) = (\alpha^{\theta} - \underline{\alpha}^{\theta})/(\overline{\alpha}_i^{\theta} - \underline{\alpha}^{\theta}) = \frac{\alpha^{\theta}}{\overline{\alpha}_i^{\theta}} \tag{18}$$

where θ is the parameter determining the shape of the distribution, $\alpha = \frac{1}{\gamma}$ the firm's marginal cost, $\overline{\alpha}_i$ the maximum marginal cost (or minimum productivity) to produce in country i, and $\underline{\alpha}$ the minimum marginal cost (or maximum productivity). Therefore, naming N the number of firms and considering solely exports from country i to country j we have:

$$x_{ij}(\alpha) = X_j(\frac{p(\alpha)}{(\sum_l N_l \int_0^{\alpha_{lj}^*} p_{lj}(\alpha)^{1-\sigma} dG(\alpha))^{\frac{1}{1-\sigma}}})^{1-\sigma} \tag{19}$$

Firms present monopolistic competition and a CES demand function. Therefore, by introducing insecurity costs – discussed in section 1 – and other sorts of trade costs, we obtain the following price and profit functions for each variety:

$$p_{ij}(\alpha) = \frac{\sigma}{\sigma - 1} w_i T_{n,ij} \alpha \tag{20}$$

$$\pi_{ij} = (\frac{x_{ij}(\alpha)}{\sigma}) - F_{n,ij} \tag{21}$$

where $T_{n,ij} = \Pi^n \tau_{n,ij}$ is a product of variable trade costs with n the n potential source of iceberg costs, including $\tau_{s,ij}$ the variable insecurity cost (derived from the expropriation risk), but also all variable trade costs sensitive to other parameters (geography, standard trade policies, institutions, etc.)., and similarly, $F_{n,ij}$ a vector of fixed trade costs that firms have to pay to enter country j from i, including $f_{s,ij}$, the fixed insecurity cost derived from insecurity barriers, but also all fixed trade costs sensitive to other sources n.

Once all exports from i to j have been aggregated, taking into account (22) and (23), and setting $\underline{\alpha} = 0$ for solving issues[66], we can define total exports from i to j:

$$X_{ij} = X_j \frac{N_i w_i^{1-\sigma} V_{ij} T_{n,ij}^{1-\sigma}}{\sum_l N_l w_l^{1-\sigma} V_{lj} T_{n,lj}^{1-\sigma}} \tag{22}$$

where V is defined as in Helpman et al. (2008):

$$V_{ij} = \int_0^{\alpha_{ij}^*} \alpha^{1-\sigma} dG(\alpha) \tag{23}$$

α_{ij}^* is by definition the level of productivity for which the profit from exporting, π_{ij}, is zero:

$$\pi_{ij} = (\frac{x_{ij}(\alpha^*)}{\sigma}) - F_{n,ij} = 0 \tag{24}$$

where:

$$\alpha^* = (\sigma - 1)\sigma^{\frac{\sigma}{\sigma-1}} (\frac{X_j}{P_l' F_{n,ij}})^{\frac{1}{\sigma-1}} \frac{1}{w_i T_{n,ij}} \tag{25}$$

with index price

$$P_l' = \sum_l N_l \int_0^{\alpha_{lj}^*} p_{lj}(\alpha)^{1-\sigma} dG(\alpha) \tag{26}$$

Once equation (27) is plugged into equation (25), we can develop V. Combined with equation (24), we obtain the following expression for bilateral exports:

$$X_{ij} = X_j \frac{N_i (\overline{\alpha}_i w_i)^{-\theta} T_{n,ij}^{-\theta} F_{n,ij}^{-[\frac{\theta}{\sigma-1}-1]}}{\sum_l N_l (\overline{\alpha}_l w_l)^{-\theta} T_{n,lj}^{-\theta} F_{n,lj}^{-[\frac{\theta}{\sigma-1}-1]}} \tag{27}$$

Finally, after defining the importer multilateral resistance term:

$$\Phi_j = (\sum_l N_l \overline{\alpha}_l^{-\theta} w_l^{-\theta} T_{n,lj}^{-\theta} F_{n,lj}^{-[\frac{\theta}{\sigma-1}-1]})^{-\frac{1}{\theta}} \tag{28}$$

we obtain the structural gravity equation:

$$X_{ij} = N_i \overline{\alpha}_i^{-\theta} w_i^{-\theta} \frac{X_j}{\Phi_j^{-\theta}} T_{n,ij}^{-\theta} F_{n,ij}^{-[\frac{\theta}{\sigma-1}-1]} \tag{29}$$

Or outlying insecurity costs from trade cost aggregates:

$$X_{ij} = N_i \overline{\alpha}_i^{-\theta} w_i^{-\theta} \frac{X_j}{\Phi_j^{-\theta}} \tau_{s,ij}^{-\theta} f_{s,ij}^{-[\frac{\theta}{\sigma-1}-1]} T_{n\neq s,ij}^{-\theta} F_{n\neq s,ij}^{-[\frac{\theta}{\sigma-1}-1]} \tag{30}$$

[66]This implies that we assume that there is always a firm that is productive enough to export at least a small amount.

with $T_{n\neq s,ij}$ and $F_{n\neq s,ij}$ respectively the variable and fixed trade costs sets excluding insecurity costs $\tau_{s,ij}$ and $f_{s,ij}$.

B.2 Extension with non-constant trade elasticities

A way to deal theoretically with the constant trade elasticity issue is to rule out the CES assumption in favour of other utility functions consistent with the idea of sub-convex gravity (Mrázová and Neary, 2017). Yet, the CES utility function has the advantage of being simple to use, works well with Pareto productivity distributions, and allows for a structural gravity equation with heterogeneous firms (Chaney, 2008). We propose another solution to capture non-constant trade elasticities: relaxing the assumption of constant technology in our standard model.

We start with the same structure as earlier, that is to say, the CES demand function, monopolistic competition and firm heterogeneity with a Pareto distribution of productivity. Yet, we allow the Pareto shape parameter θ to be exporter specific and assume that worldwide technology has been developed to match the sub-variety of dominant markets. Precisely, we apply a distortion ζ_j on the productivity distribution. In keeping with empirical evidences presented in Carrère et al. (2020) and section 3.5, we assume $\frac{\delta\theta_i}{\delta\frac{Y_i}{\Pi_i^{-\theta_i}}} < 0$ and $\frac{\delta\zeta_j}{\delta\frac{X_j}{\Phi_j^{-\theta_j}}} < 0$. The larger countries i's and j's share of global revenue, the greater the firm's probability of dropping high productivity. In other words, firms still have a monopoly on their variety, but produce a sub-variety for each destination that has its own productivity level and a greater chance of being more productive in sub-varieties designed for the largest markets. For example, a firm might produce blue T-shirts (its variety) with the highest level of productivity for the sub-variety designed for the US (chemical dyed large blue T-shirts). However, it would have to divert from this optimal production chain when producing the sub-variety for Sweden (naturally dyed long blue T-shirts), where consumer characteristics and preferences are different.

We define the country-pair specific productivity distribution as follows:

$$G_{ij}(\alpha) = (\alpha^{\zeta_j\theta_i} - \underline{\alpha}^{\zeta_j\theta_i})/(\overline{\alpha}_i^{\zeta_j\theta_i} - \underline{\alpha}^{\zeta_j\theta_i}) = \frac{\alpha^{\zeta_j\theta_i}}{\overline{\alpha}_i^{\zeta_j\theta_i}} \quad (31)$$

Therefore, the firms' probability of dropping sub-variety productivity φ_{ij} and the productivity distribution in the economy are determined by θ_i and ζ_j. Each country-pair presents a specific productivity distribution, while each firm drops different outputs depending on the destination market. Nonetheless, because each country-pair's productivity is Pareto distributed, the observed average productivity of firms in a given origin country is also Pareto distributed.[67] In this extended model, the price

[67] At country-pair level, many firms drop small productivity while a few firms drop high productivity. Taking the firms' average productivity, we observe the same thing. Many firms drop small productivity for all destinations while a few firms drop only high productivity. Nonetheless, in this framework, some firms that are on average not very productive may have a low level of productivity for the large majority of destinations, but a high level of productivity for a few

and profit equations are unchanged. Yet, because the productivity distribution is different from the standard model, the firms' productivity is affected and their prices too.

Running through the model, we obtain the following expression of the structural gravity equation with non-constant trade cost elasticities:

$$X_{ij} = N_i \overline{\alpha}_i^{-(\zeta_j \theta_i)} w_i^{-(\zeta_j \theta_i)} \frac{X_j}{\Phi'_j} \tau_{s,ij}^{-(\zeta_j \theta_i)} f_{s,ij}^{-[\frac{\zeta_j \theta_i}{\sigma-1}-1]} T_{n \neq s,ij}^{-(\zeta_j \theta_i)} F_{n \neq s,ij}^{-[\frac{\zeta_j \theta_i}{\sigma-1}-1]} \tag{32}$$

with

$$\Phi'_j = \sum_l N_l \overline{\alpha}_l^{-\zeta_j \theta_l} w_l^{-\zeta_j \theta_l} T_{n,lj}^{-\zeta_j \theta_l} F_{n,lj}^{-[\frac{\zeta_j \theta_l}{\sigma-1}-1]} \tag{33}$$

This distorted structural gravity equation raises new trade cost elasticities: variable trade cost elasticity $-(\zeta_j \theta_i)$ and fixed trade cost elasticity $-[\frac{\zeta_j \theta_i}{\sigma-1}-1]$. The lower θ_i and ζ_j, the lower the trade elasticities. Therefore, large economies are less affected by trade costs. This explains why we observe trade cost elasticities inversely proportional to the value of trade (cf. section 3.5).

C Two-way (robust) fixed effects estimations

As pointed up by De Chaisemartin and d'Haultfoeuille (2020), heterogeneous treatments or treatment effects over time and groups may return false results in the case of two-way fixed effects estimations. Comparing groups (here country-pairs) that are not treated at the same time or that experience different outcomes following the treatment could cause negative weights in the (bias) ATE. The estimator developed by De Chaisemartin and d'Haultfoeuille (2020) is sufficient to deal with this bias. But, because we want to address the effect's dynamic, we use the later estimator developed by De Chaisemartin and D'Haultfoeuille (2022), which provides event study results.

The intuition behind this estimation is that avoiding these negative weights entails a comparison of the first-time switchers' t-1 to t+l outcome evolution with the t-1 to t+l outcome evolution of country-pairs whose treatment has hitherto remained stable; with t the treatment time and l the event time. Our panel is balanced, so we estimate the treatment effect from positive switchers (i.e. alliance signatures):

$$\delta_l = \frac{DID_{+,l}}{DID^D_{+,l}} \tag{34}$$

where:

$$DID_{+,l} = \sum_{g:D_{g,1}=0, F_g<T_u-l} \frac{N_{g,t_g+l} \beta^{t_g+l}}{N^1_l} \left[(Y_{g,t_g+l} - Y_{g,t_g-1}) - \sum_{g:D_{g',1}=0, t_{g'} t_{g'} > t_g+l} \frac{N_{g',t+l}}{N^u_{t_g+l}} (Y_{g',t_g+l} - Y_{g',t-1}) \right] \tag{35}$$

destinations.

and

$$DID^D_{+,l} = \sum_{g:D_{g,1}=0,F_g<T_u-l} \frac{N_{g,t_g+l}\beta^{t_g+l}}{N^1_l} \left[(D_{g,t_g+l} - D_{g,t_g-1}) - \sum_{g:D_{g',1}=0,t_{g'}t_{g'}>t_g+l} \frac{N_{g',t+l}}{N^u_{t_g+l}} (D_{g',t_g+l} - D_{g',t-1}) \right] \tag{36}$$

with t_g the time of group g's treatment variation, $\beta \in (0,1]$ the planner's discount rate, D the treatment variation, T_u the last observed period with a group untreated since period 1, $N^1_l = \sum_{g:D_{g,1}=0,t_g<T_u-l} N_{g,t_g+l}\beta^{t_g+l}$ the discounted number of units in groups reaching l periods after their first treatment or before T_u, and Y the outcome – in our case exports. Yet, since ALL_{ijt} is a dummy variable, $DID^D_{+,l} = 1$. Therefore, we have:

$$\delta_l = DID_{+,l} \tag{37}$$

We present the results for military alliances in table 8 and defence pacts in table 9. They show a growing dynamic effect and confirm the robustness of our identification strategy.

Table 8: Military alliances: Chaisemartin D'Haultfoeuille (2020) estimator

Dependent variable: Bilateral exports						
Time	t	t+1	t+2	t+3	t+4	t+5
Alliance's coeff	0.339^a	0.457^a	0.500^a	0.581^a	0.661^a	0.682^a
Standard error	(0.085)	(0.086)	(0.090)	(0.093)	(0.093)	(0.094)
RTA control	yes	yes	yes	yes	yes	yes
No. observ.	128.004	127.878	127.788	127.594	121.022	114.458
No. switchers	250	248	248	246	242	238

Note: Dependent variable is the inverse hyperbolic sine transformation of exports from country i to country j at time t in millions of current dollars. t is the year of the pair's alliance signature. Standard errors clustered at country-pair level are in parentheses. a, b and c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 9: Defence pacts: Chaisemartin D'Haultfoeuille (2020) estimator

Dependent variable: Bilateral exports						
Time	t	t+1	t+2	t+3	t+4	t+5
Alliance's coeff	0.434^a	0.561^a	0.568^a	0.656^a	0.714^a	0.748^a
Standard error	(0.118)	(0.119)	(0.118)	(0.123)	(0.127)	(0.128)
RTA control	yes	yes	yes	yes	yes	yes
No. observ.	44.888	44.854	44.844	44.818	38.266	38.262
No. switchers	172	172	172	172	170	170

Note: Dependent variable is the inverse hyperbolic sine transformation of exports from country i to country j at time t in millions of current dollars. t is the year of the pair's alliance signature. Standard errors clustered at country-pair level are in parentheses. a, b and c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

D The Cold War

How to define the Cold War and to what extent it can bias our estimation is not obvious. If we consider that the Cold War is a period structuring global country relationships, it is a time-level variable and is captured by our country-year fixed effects. If we define the Cold War as a period when countries were either capitalist or communist, it comes under the country-year level and once again is captured by our fixed effects. Yet, if we consider the Cold War as a latent conflict between the western bloc and the eastern bloc, our fixed effects are not sufficient. Therefore, we include in our estimation a dummy variable taking the value one if country i is a member of one bloc and country j a member of the opposite bloc, conditional on the absence of military alliances between i and j, and zero otherwise.[68]

In this case, we deal with highly specific heterogeneity in alliance effects. The Cold War variable is a particular case of a non-alliance relationship between countries i and j. In section 3.3, the defence pact coefficient is estimated compared with the average case of non-alliances. Yet, introducing the Cold War variable would exclude a case of "latent-conflict-non-alliance" from this average. Therefore, the non-alliance average would be closer to a neutral relationship. This could induce overestimated coefficients. estimation should address this issue. Using propensity score matching, the treated and control groups should be comparable in terms of bilateral diplomatic relationships. Controlling for the $Cold.war_{ijt}$ dummy, we estimate the defence pacts effect as in section 3.2 and with a DATT. The results are reported in table 10. In the standard estimation (column 1), controlling for the Cold War reduces the defence pact coefficient. In the DATT estimation (column 2), variables of interest coefficients are not impacted while the Cold War's coefficient is non-significant. In both estimations, the defence pact effect on bilateral exports is estimated at 80% and is highly significant.

[68] $Cold.war_{ijt} = \mathbb{1}_{\{alliance_{ijt}=0\}} * ((alliance_{i,USA,t} + \mathbb{1}_{\{i=USA\}}) * (alliance_{RUS,jt} + \mathbb{1}_{\{j=RUS\}}) + (alliance_{i,RUS,t} + \mathbb{1}_{\{i=RUS\}}) * (alliance_{USA,jt} + \mathbb{1}_{\{j=USA\}})) * \mathbb{1}_{\{alliance_{i,USA,t} * alliance_{USA,jt}=0\}} * \mathbb{1}_{\{alliance_{i,RUS,t} * alliance_{RUS,jt}=0\}}$

Table 10: Defence pacts and the Cold War

Estimator: PPML Dependent variable: Exports Variables	Standard three-way FE	DATT
Defence pact	0.592^a	
	(0.078)	
t-4		-0.037
		(0.118)
t-3		0.063
		(0.077)
t-2		-0.047
		(0.086)
t-1		0.009
		(0.087)
t		0.242^a
		(0.088)
t+1		0.378^a
		(0.085)
t+2		0.420^a
		(0.079)
t+3		0.515^a
		(0.077)
t+4		0.614^a
		(0.072)
>=t+5		0.574^a
		(0.042)
Weak Alliance	-0.127	-0.463^a
	(0.159)	(0.105)
Cold War	-0.239^a	-0.028
	(0.076)	(0.031)
RTA	0.135^a	0.156^a
	(0.032)	(0.014)
Exporter x Year FE	yes	yes
Importer x Year FE	yes	yes
Dyadic FE	yes	yes
No. observ.	320.666	39,488

Note: PPML, Poisson Pseudo Maximum Likelihood; FE, Fixed effects; DATT, Differenced Average Treatment on the Treated; Dependent variable is exports from country i to country j at time t. Robust standard errors are in parentheses; column (1) standard errors are clustered at country-pair level; column (2) standard errors are clustered at country-pair-year level. a, b and c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

E Supplementary tables and figures

Table 11: List of countries

Albania	Denmark	Latvia	Saudi Arabia
Algeria	Ecuador	Kyrgyzstan	Russia
Argentina	Egypt	Libya	Serbia-Montenegro
Australia	Estonia	Lithuania	Singapore
Austria	Finland	Luxembourg	Slovakia
Bangladesh	France	Malaysia	Slovenia
Belarus	Gabon	Malta	South Korea
Belgium	Germany	Mexico	Spain
Bolivia	Greece	Morocco	Sri Lanka
Bosnia and Herzegovina	Hong Kong	Netherlands	Sweden
Brazil	Hungary	New Zealand	Switzerland
Brunei	Iceland	Nigeria	Taiwan
Bulgaria	India	North Macedonia	Thailand
Cameroon	Indonesia	Norway	Tunisia
Canada	Ireland	Pakistan	Turkey
Chile	Israel	Paraguay	Ukraine
China	Italy	Peru	United Kingdom
Columbia	Ivory Coast	Philippines	United States of America
Croatia	Japan	Poland	Uruguay
Cyprus	Kazakhstan	Portugal	Venezuela
Czech Republic	Kenya	Romania	Vietnam

Table 12: Exports and military alliances, robustness checks

Robustness check:	Intra. trade	Tariffs			Tariffs & Intra. trade	RTA depth		Geo. and eco. dist.	Extended panel	Import shares
Period:	1989-2012	1996-2012			1996-2012	1967-2012		1967-2012	1948-2012	1967-2012
Baseline sample:	No	Yes	No	No	No	No		Yes	No	Yes
Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Alliance	0.490^{a}	0.203^{a}	0.157^{b}	0.159^{b}	0.148^{b}	0.439^{a}	0.438^{a}	0.473^{a}	0.199^{a}	0.646^{a}
	(0.046)	(0.052)	(0.067)	(0.068)	(0.071)	(0.116)	(0.116)	(0.106)	(0.046)	(0.105)
RTA	0.104^{a}						-0.021	0.157^{a}	0.043	0.374^{a}
	(0.041)						(0.052)	(0.033)	(0.049)	(0.045)
ln(Tariffs+1)				-0.492	-1.987^{a}					
				(0.335)	(0.432)					
Intra x Year	-0.039^{a}				-0.026^{a}					
	(0.003)				(0.006)					
RTA depth						0.010^{a}	0.011^{a}			
						(0.001)	(0.002)			
Diff. GDP per cap.								-0.005		
								(0.004)		
ln(Distance) x Year								-0.000		
								(0.001)		
Exporter x Year FE	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes
Importer x Year FE	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes
Dyadic FE	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes
No. observ.	97.129	118.439	54.035	54.035	36.664	314.128	314.128	320.666	773.689	320.666

Note: PPML, Poisson Pseudo Maximum Likelihood; FE, Fixed effects; Dependent variable is exports from country i to country j at time t. Standard errors clustered at the country-pair level are in parentheses. Because of data limitation, several estimations are performed with a shorter panel or omitted observations/country-pairs.
In column 2 we perform the same regression as in column 1 but without observations whose Bilateral average tariff is unknown.
a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 13: Corrected three-way ppml estimation

Estimator: PPML Dependent variable: exports	
Variables	(1)
Alliance	0.482^a
Corrected bias	-0.010
	(0.118)
RTA	0.166^a
Corrected bias	-0.005
	(0.036)
Exporter x Year FE	yes
Importer x Year FE	yes
Dyadic FE	yes
No. observ.	320.666

Note: PPML, Poisson Pseudo Maximum Likelihood; FE, Fixed effects; Dependent variable is exports from country i to country j at time t in millions of current dollars. Standard errors clustered at country-pair level are in parentheses. Coefficients are corrected from the asymptotic bias.
a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 14: Extensive and intensive margins of trade

A: Intensive margin

Dependent variable:	$\{X_{ijt} \mid X_{ijt} > 0, X_{ijt-1} > 0\}$
Estimator:	PPML
Alliance	0.389^a
	(0.094)
Controls	yes
Exporter x Year FE	yes
Importer x Year FE	yes
Dyadic FE	yes
No. observ.	289.581

B: Extensive margin

Dependent variable:	Export dummy
Estimator:	Logit
Alliance	0.403^a
	(0.079)
Controls	yes
Year FE	yes
Dyadic FE	yes
No. observ.	102.442

Note: PPML, Poisson Pseudo Maximum Likelihood; FE, Fixed effects; Dependent variable is exports from country i to country j at time t without the zero observations and conditional to a positive value the previous year. Standard errors clustered at country-pair level are in parentheses.
Intensive margin estimation's control: RTAs; Extensive margin estimation's controls: RTAs, exporter's GDP, importer's GDP, exporter's population, importer's population. Coefficient of our Logit estimation can be directly interpreted as the elasticity.
a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 15: Exports and defence pacts

Estimator: PPML		
Dependent variable: Bilateral exports		
Variables	(1)	(2)
Defence pact	0.693[a]	0.694[a]
	(0.079)	(0.120)
Weak alliance	-0.107	-0.149
	(0.151)	(0.153)
RTA	0.137[a]	0.137[a]
	(0.032)	(0.032)
NATO		0.643[a]
		(0.080)
Exporter x Year FE	yes	yes
Importer x Year FE	yes	yes
Dyadic FE	yes	yes
No. observ.	320.666	320.666

Note: PPML, Poisson Pseudo Maximum Likelihood; FE, Fixed effects; Dependent variable is exports from country i to country j at time t in millions of current dollars. Standard errors clustered at country-pair level are in parentheses.
a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Figure 7: Quantile estimates of RTAs

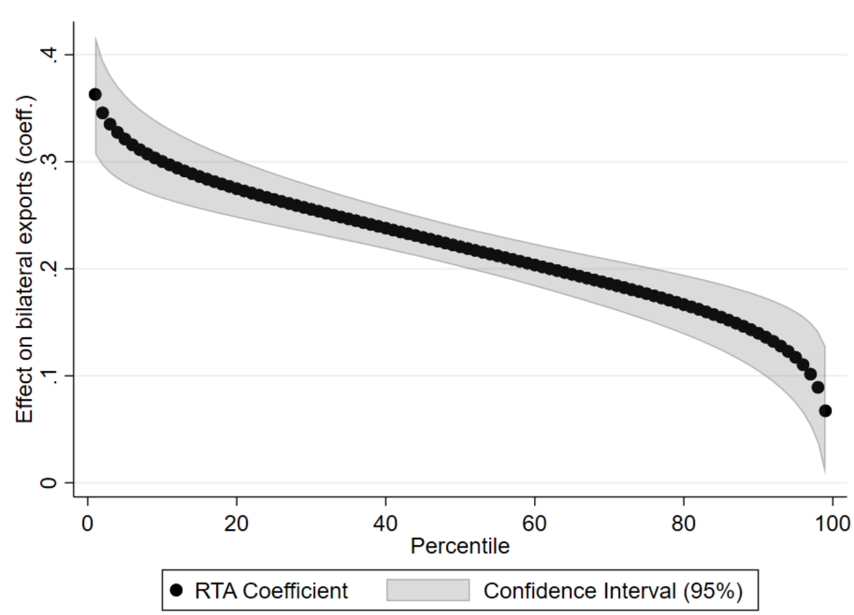


Table 16: Alliances and bilateral exports, IV control function

Dependent variable: exports		
Second stage	(1)	(2)
Estimator:	PPML	PPML
Alliance	0.635^a	
	(0.025)	
Defence pact		0.624^a
		(0.025)
Weak alliance		-0.011^a
		(0.041)
RTA	0.152^a	0.144^a
	(0.012)	(0.011)
First stage residuals	-0.357^a	0.176^a
	(0.049)	(0.057)
Exporter x Year FE	yes	yes
Importer x Year FE	yes	yes
Dyadic FE	yes	yes
No. observ.	320.666	320.666
First-stage		
Estimator:	OLS	OLS
Dependent variable:	Alliance	Defence pact
Common out. alliances	0.056^a	0.053^a
	(0.007)	(0.006)
Weak alliance		-0.398^a
		(0.062)
RTA	0.011	0.015^c
	(0.008)	(0.009)
Exporter x Year FE	yes	yes
Importer x Year FE	yes	yes
Dyadic FE	yes	yes
No. observ.	320.712	320.712
KPW F-stat	67	78
KPW LM-stat	11	10

Note: OLS, Ordinary Least Squares; FE, Fixed effects. Dependent variable is exports from country i to country j at time t in millions of current dollars. Standard errors clustered at the exporter and importer levels are in parentheses. Second-stage standard errors are bootstrapped.
a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 17: Outside alliances

Estimator: PPML		
Dependent variable: Exports		
Variables	(1)	
Alliance	0.467^{a}	
	(0.108)	
Outside alliance(exp.)	-0.030	
	(0.174)	
Outside alliance(imp.)	-0.106	
	(0.157)	
Defence pact		0.702^{a}
		(0.081)
Weak alliance		-0.113
		(0.169)
Outside defence pact(exp.)		0.101
		(0.152)
Outside defence pact(imp.)		0.044
		(0.144)
Outside weak alliance(exp.)		-0.102
		(0.185)
Outside weak alliance(imp.)		0.101
		(0.266)
RTA	0.160^{a}	0.131^{a}
	(0.033)	(0.032)
Exporter x Year FE	yes	yes
Importer x Year FE	yes	yes
Dyadic FE	yes	yes
No. observ.	320.666	320.666

Note: PPML, Poisson Pseudo Maximum Likelihood; FE, Fixed effects. Dependent variable is exports from country i to country j at time t. Standard errors clustered at country-pair level are in parentheses. a, b and c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 18: Alliances and bilateral exports, 2SLS

Dependent variable: exports		
Second stage		
	(1)	(2)
Estimator:	OLS	2SLS
Instrument variable:	None	Common out. alliances
Alliance	0.664^a	0.452^a
	(0.088)	(0.165)
RTA	0.562^a	0.575^a
	(0.029)	(0.074)
Exporter x Year FE	yes	yes
Importer x Year FE	yes	yes
Dyadic FE	yes	yes
No. observ.	320.712	320.712
First-stage		
Instrumented variable:	None	Alliance
Common out. alliances		0.056^a
		(0.007)
RTA		0.012^a
		(0.008)
KPW rk F-stat:		67
KPW rk LM-stat:		11

Note: OLS, Ordinary Least Squares; 2SLS, Two-Stage Least Square; FE, Fixed effects; Dependent variable is the inverse hyperbolic sine transformation of exports from country i to country j at time t in millions of current dollars. Standard errors clustered at country-pair levels (column 1) and at the importer and exporter levels (column 2) are in parentheses
a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 19: Dynamic propensity score matching, some details

Estimator: OLS Dependent variable: Defence pact			
Variables	(1)	(2)	(3)
Year:	1967	1990	2012
Exp. ln(GDP)	0.379^a	0.314^a	0.318^a
	(0.031)	(0.025)	(0.027)
Imp. ln(GDP)	0.379^a	0.314^a	0.318^a
	(0.031)	(0.025)	(0.027)
Exp. ln(Pop.)	-0.185^a	-0.117^a	-0.116^a
	(0.035)	(0.029)	(0.028)
Imp. ln(Pop.)	-0.185^a	-0.117^a	-0.116^a
	(0.035)	(0.029)	(0.028)
Common religion	1.231^a	1.616^a	1.378^a
	(0.098)	(0.098)	(0.091)
ln(Distance)	-0.435^a	-0.331^a	-0.503^a
	(0.031)	(0.033)	(0.030)
Common official language	1.169^a	1.023^a	0.817^a
	(0.078)	(0.082)	(0.079)
Colonial past	-0.479^a	-0.641^a	-0.366^a
	(0.140)	(0.154)	(0.134)
No. observ.	6.970	6.870	6.866

Note: OLS, Ordinary Least Square; Dependent variable is the presence of a military alliance between countries i and j at time t. Standard errors are in parentheses. The PSM is made for each year from 1967 to 2012, respectively.
a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 20: Propensity score matching, variables' mean

Variable	Psm base	Standard base
GDP exp. (ln)	11.633	10.608
GDP imp. (ln)	11.629	10.608
Population exp. (ln)	2.816	2.575
Population imp. (ln)	2.812	2.575
Distance (ln)	7.762	8.511

Table 21: DATT

Estimator: PPML Dependent variable: Bilateral exports	
Variables	(1)
t-4	0.026
	(0.116)
t-3	0.074
	(0.075)
t-2	-0.036
	(0.085)
t-1	0.020
	(0.086)
t	0.253[a]
	(0.087)
t+1	0.389[a]
	(0.083)
t+2	0.430[a]
	(0.078)
t+3	0.526[a]
	(0.076)
t+4	0.625[a]
	(0.070)
>=t+5	0.584[a]
	(0.039)
Weak Alliance	-0.455[a]
	(0.107)
RTA	0.156[a]
	(0.014)
Exporter x Year FE	yes
Importer x Year FE	yes
Dyadic FE	yes
No. observ.	39.488

Note: PPML, Poisson Pseudo Maximum Likelihood; FE, Fixed effects; Year is standardised with Year=1 in 1967. Dependent variable is exports from country i to country j at time t in millions of current dollars. t is the date of the signatory of the alliance between country i and j. Alliance's signatory effect is estimated in comparison to $k' <= t-5$. Standard errors clustered at country-pair-year level are in parentheses. Observations are weighted in function of our Propensity Score Matching.
a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 22: The military cooperation channel

Second stage	(1)	(2)
Estimator:	PPML	
Dependent variable:	Exports	
Military cooperation	0.401^{a}	0.294^{a}
	(0.069)	(0.044)
Weak alliance	-0.227	-0.286
	(0.157)	(0.178)
RTA	0.040	0.041
	(0.036)	(0.040)
Exporter x Year FE	yes	yes
Importer x Year FE	yes	yes
Dyadic FE	yes	yes
No. observ.	167.304	167.304
First-stage		
Estimator:	OLS	
Dependent variable:	Military cooperation	
Defence pact	1.004^{a}	
	(0.100)	
Common out. def. pacts		0.081^{a}
		(0.005)
Weak alliances	0.241^{b}	0.178^{c}
	(0.097)	(0.096)
RTA	0.128^{a}	0.125^{a}
	(0.035)	(0.035)
Exporter x Year FE	yes	yes
Importer x Year FE	yes	yes
Dyadic FE	yes	yes
No. observ.	167.328	167.328

Note: OLS, Ordinary Least Square; PPML, Poisson Pseudo Maximum Likelihood; FE, Fixed effects. The panel starts in 1989. Common defence pacts sum all external partners for which country i and j both have a defence pact. Military cooperation is the inverse hyperbolic sine transformation of $\sum coop.mil.ev._{ijt}$. Robust standard errors clustered at country-pair level are in parentheses. Second-stage standard errors are bootstrapped.
a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 23: Estimating the trade elasticity

Estimator: PPML Dependent variable: Bilateral exports	
Variables	(1)
ln(1+Tariffs)	-3.754^{a}
	(1.102)
ln(Distance)	-0.781^{a}
	(0.044)
Common religion	-0.157
	(0.160)
Contiguity	0.579^{a}
	(0.111)
Common language	0.131
	(0.101)
Exporter x Year FE	yes
Importer x Year FE	yes
Dyadic FE	no
No. observ.	54,479

Note: PPML, Poisson Pseudo Maximum Likelihood; FE, Fixed effects; Dependent variable is exports from country i to country j at time t in millions of current dollars. Standard errors clustered at country-pair level are in parentheses.
a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 24: GE Exports and Welfare with heterogeneous elasticities

Country (Iso3)	Exports	Real revenues	Country (Iso3)	Exports	Real revenues
ALB	2.53	0.20	ISL	26.34	9.52
ARG	13.55	4.07	ISR	-1.21	-1.04
AUS	2.64	0.26	ITA	20.32	2.85
AUT	-0.41	-1.25	JPN	5.10	0.32
BEL	11.47	11.80	KEN	-2.90	-0.17
BGD	0.32	-0.55	KOR	3.34	0.40
BGR	7.45	1.25	LKA	-0.77	-0.48
BOL	22.29	6.68	MAR	1.19	-0.15
BRA	12.81	0.57	MEX	16.47	7.03
CAN	24.08	6.14	MLT	1.01	-1.30
CHE	-0.16	-1.30	MYS	0.65	-0.49
CHL	8.99	3.93	NGA	1.30	-1.68
CHN	-0.33	-0.11	NLD	10.41	12.05
CIV	10.91	0.21	NOR	12.95	4.32
CMR	0.77	-0.54	NZL	-0.07	-0.26
COL	23.47	1.89	PAK	16.31	3.20
CYP	-0.33	-1.20	PER	17.16	1.70
DEU	15.54	3.87	PHL	3.21	1.07
DNK	17.04	6.10	POL	34.14	3.22
DZA	0.40	-0.50	PRT	33.81	4.53
ECU	21.34	3.90	ROM	2.66	0.04
EGY	-1.85	-0.38	SAU	0.58	-0.60
ESP	25.75	3.11	SER	-3.29	-0.27
FIN	-1.11	-0.78	SGP	-0.04	-0.24
FRA	24.91	4.31	SWE	-0.97	-1.47
GAB	1.31	3.06	THA	0.42	-0.44
GBR	22.94	4.09	TUN	1.56	-0.22
GRC	36.39	3.59	TUR	35.42	1.75
HKG	0.28	-0.71	URY	29.96	2.76
HUN	17.40	7.04	USA	34.09	2.02
IDN	0.32	-0.21	VEN	11.75	5.72
IND	-1.13	-0.15	VNM	0.36	-0.29
IRL	1.35	-2.07			

Note: The real revenue is our measure of welfare. All numbers are variations in percentage.

Table 25: Alliances scenarios and Welfare

Scenarios:	NATO expansion	NATO East-rupture	New East-block	Scenarios:	NATO expansion	NATO rupture	New East block
Country (Iso3)	Real revenues	Real revenues	Real revenues	Country (Iso3)	Real revenues	Real revenues	Real revenues
ALB	23.81	-1.24	6.24	ISL	1.60	-0.37	-0.45
ARG	-0.20	0.04	0.01	ISR	-0.69	0.12	0.12
AUS	3.29	0.02	0.02	ITA	1.08	-0.27	-0.33
AUT	17.46	0.40	0.20	JPN	1.61	0.01	-0.00
BEL	1.80	-0.38	-0.46	KEN	-0.17	0.01	-0.02
BGD	-0.49	0.07	-0.08	KOR	3.07	0.03	0.01
BGR	10.27	-2.67	2.52	LKA	-0.39	0.03	-0.03
BOL	-0.15	0.00	-0.02	MAR	-0.39	0.09	0.05
BRA	-0.07	0.01	0.01	MEX	-0.25	0.04	-0.01
CAN	1.28	-0.02	-0.03	MLT	23.44	0.25	0.20
CHE	17.13	0.19	0.15	MYS	-0.54	0.04	0.00
CHL	-0.46	0.06	0.02	NGA	-1.03	0.12	0.06
CHN	-0.07	0.01	0.09	NLD	2.27	-0.50	-0.65
CIV	-0.06	0.02	0.01	NOR	2.63	-0.23	-0.42
CMR	-0.38	0.06	0.03	NZL	6.47	0.02	0.00
COL	2.05	0.02	0.00	PAK	6.83	0.07	0.81
CYP	-1.24	0.32	0.25	PER	-0.24	0.01	-0.00
DEU	1.88	-0.59	-0.67	PHL	-0.46	0.03	0.02
DNK	3.40	-0.46	-0.72	POL	0.73	-5.60	-3.50
DZA	-0.38	0.09	0.07	PRT	0.57	-0.10	-0.12
ECU	-0.26	0.03	-0.01	ROM	11.20	-0.65	3.98
EGY	-0.33	0.07	0.02	SAU	-0.71	0.08	0.09
ESP	0.68	-0.19	-0.22	SER	6.52	0.24	3.25
FIN	11.51	0.13	4.43	SGP	-0.27	0.01	0.02
FRA	1.00	-0.22	-0.25	SWE	15.73	0.19	3.38
GAB	-0.82	0.16	0.29	THA	-0.54	0.04	0.01
GBR	1.71	-0.22	-0.26	TUN	-0.53	0.12	0.09
GRC	0.96	-1.04	-1.25	TUR	0.40	-2.73	-1.64
HKG	-1.14	0.07	0.18	URY	-0.15	0.03	0.01
HUN	2.93	-12.05	-5.57	USA	0.03	-0.03	-0.04
IDN	-0.32	0.02	-0.00	VEN	-0.31	0.03	0.03
IND	-0.13	0.02	0.00	VNM	-0.33	0.03	-0.03
IRL	24.58	0.18	0.11				

Note: The real revenue is our measure of welfare. All numbers are variations in percentage.