Product substitutability and the Distance Puzzle*

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Abstract

We propose a novel explanation of the non-decreasing distance elasticity of trade over 1962-2013, commonly referred to as 'the distance puzzle'. Consistently with previous work, we find that the changing composition of world trade has contributed to making trade less rather than more sensitive to distance. Our original contribution consists in documenting a 35% increase in the perceived substitutability of products traded on the world market over the same period. We find that the Armington elasticity of substitution increased from about 1.7 to 2.3 between 1965 and 2013. This elasticity corresponds to the elasticity of trade flows to trade costs on the intensive margin. In the Armington framework it determines the elasticity of trade flows to trade costs. The evolution of this parameter suffices to explain the non-decreasing distance elasticity of trade.

Keywords: gravity equation, distance puzzle, trade elasticity, trade costs

JEL codes: F15, N70

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Introduction

It is a well-established fact that trade has not become less sensitive to distance over the last 60 years. Disdier and Head (2008) adopt a meta-analytical approach and find that distance impedes trade by 37% more in the 1990s than it did from 1870 to 1969[1]. Head and Mayer (2013) estimate the distance elasticity of trade in successive cross-sections and find that it has doubled in 1960-2005. This increase in the distance elasticity of trade has been dubbed the "distance puzzle", as the common opinion is that technological developments in transportation and communication, e.g. the airplane, the container, and the internet, would have led to the "death of distance" by the end of the 20$^{\text{th}}$ century[2].

In this paper we investigate the empirical relevance of a different mechanism that may help to rationalize the distance puzzle. We make the simple point that the flattening out of the world may go hand in hand with a persistent impeding effect of distance on trade if consumers perceive product bundles shipped out by each country to the world market as increasingly substitutable.[3] We work with the canonical Anderson and van Wincoop (2003) framework to quantify the contribution of this mechanism. We find that the non-decreasing distance elasticity of trade can be explained by the increasing sensitivity of consumers to price differences.

Prominent explanations of the distance puzzle

Recent work rationalizes the distance puzzle in three complementary ways: by pointing out a possible misspecification of the econometric model, by refining the specification of the trade cost function, and, more recently, through the lens of network analysis.

The first strand of the literature investigates the incidence of the estimation method on the distance puzzle. Silva and Tenreyro (2006) advocate estimating the gravity model in multiplicative form using a specific non-linear estimator, the Poisson Pseudo Maximum Likelihood (PPML). Contrary to the canonical loglinear approach, this estimator provides consistent estimates and is robust to rounding error and overdispersion which are both likely features of trade data[4]. The magnitude of the distance puzzle is reduced when the gravity model is estimated in multiplicative form (Bosquet and Boulhol (2015); Head and Mayer (2013)).

[1] See also Berthelon and Freund (2008); Combes et al. (2006); Brun et al. (2005); Buch et al. (2004).
[2] Cairncross (1997); Levinson (2006); Friedman (2007)
[3] Equation (4) in Head and Mayer (2013): the distance effect is a product of two elasticities, the elasticity of trade to trade costs and the elasticity of trade costs to distance. We investigate the evolution of the former.
[4] Santos Silva and Tenreyro (2011) and Fally (2015) provide evidence on desirable properties of the PPML. Head and Mayer (2014) review properties of alternative estimators. See also Bosquet and Boulhol (2015, 2014).

The sensitivity of the distance puzzle to the estimation method is likely due to sample composition effects (Head and Mayer (2013); Larch et al. (2016)). Indeed, the growth of trade has been both intensive in the sense that the volume of established trade relations has increased and extensive in the sense that new trade relations have been established (Helpman et al. (2008); Baldwin and Harrigan (2011)). If trade relations have in priority been established between small and distant partners, the reduction in the number of zeros may have gradually reduced the underestimation of the distance coefficient in the loglinear specification[5]. This explanation echoes Felbermayr and Kohler (2006) who pointed out that the log-linear specification was subject to sample selection bias due to the exclusion of zero trade flows. They conjectured that the distance puzzle was an artefact of reduction in this bias through the extensive margin of trade[6].

The second and most prominent strand of the literature singles out the underpinnings of the trade cost function as key to understanding the distance puzzle. The basic point formulated by Buch et al. (2004) is that the distance elasticity of trade is invariant to reductions in transportation and communication costs if their distribution over distance remains unchanged[7]. Furthermore, while the distance elasticity of transport costs may have decreased (Hummels (2007)), other cost components, such as delays, may have become more distance-elastic (Hummels and Schaur (2013)). More generally, as argued by Head and Mayer (2013), if freight costs account for an ever smaller fraction of distance-dependent trade costs, the distance elasticity of trade will be determined by other, possibly persistent, cost components[8].

A complementary mechanism is proposed in Krautheim (2012) in the heterogeneous firms' framework. He models the informational component of trade costs as a fixed cost which decreases in the number of exporting firms. This refinement of the trade cost function magnifies the distance elasticity of trade because the number of exporters is decreasing in variable trade costs which increase with distance. This magnification mechanism may have been reinforced by the increasing weight of information costs in total fixed costs.

[5] Larch et al. (2016) attribute the puzzle to the growing bias of the OLS estimates. Using a nonlinear estimator and controlling for the number of exporting firms, they find a decreasing distance elasticity in 1980-2006.

[6] This leaves open the question of the estimator which correctly captures the level of the distance elasticity. Head and Mayer (2013) argue that PPML is giving little weight to small trade flows characteristic of more distant partners. For Silva and Tenreyro (2006) small trade flows are more prone to measurement error.

[7] Egger (2008) argues that the marginal effect of distance on trade is fundamentally non uniform across trade partners and decreasing in the level of bilateral trade. This reinstates increased openness as a possible explanatory factor because a uniform reduction in trade costs has non uniform effects on bilateral trade.

[8] Head and Mayer (2013) propose a typology of persistent but unobserved trade costs. Daudin (2003, 2005) put forward that trade costs may have remained stable as a share of value added.

An alternative explanation put forward in models with input-output linkages is that the relationship between total trade costs and transport costs may be non-monotonic. An increasing distance elasticity may be an endogenous outcome of transport cost reductions if they engender a reoptimization of the production process which ends up increasing the relative cost of long-distance trade. One possible mechanism is trade cost magnification through multiple border crossings by goods as a consequence of increased production fragmentation (Yi (2010); Daudin et al. (2011); Noguera (2012)). Another mechanism formalized by Duranton and Storper (2008) works through quality upgrading. Lower transport costs shift trade towards higher-quality inputs which are more distance-sensitive because their customization requires intensive communication, e.g. more back-and-forth travelling, between upstream and downstream firms.

A shift of focus to the trade elasticity

The focus of the literature on the shape of the trade cost function mirrors the expectation that the distance coefficient moves together with the elasticity of trade costs to distance. But Chaney (2016) provides a theoretical foundation for the gravity equation through the lens of network analysis which demonstrates that the distance coefficient can be invariant to the trade cost function. In Chaney (2016) the rate of distance decay in aggregate trade is linked to the rate of decay in the density of firms which cover that distance with their network of contacts. As the geographic dispersion of the network is increasing in firm size, the shape parameter of the firm size distribution plays a key role in explaining movements in the distance coefficient. Thus, technological advances in transportation increase the geographic dispersion of exports at the level of the firm but have no incidence on the distance elasticity of aggregate trade as long as the stationary firm size distribution verifies Zipf's law.

The link between the distance coefficient and the parameter which captures the degree of structural heterogeneity in the economy is not specific to Chaney (2016). Every theoretical foundation of the gravity model delivers a functional relationship of the distance elasticity with the intensity of the incentive to trade, e.g. the degree of structural heterogeneity in some model-specific dimension. The combination of empirical evidence on the changing shape of the trade cost function with evidence on the stability of the distance distribution of trade indicates that structural heterogeneity may have contributed to the evolution of the distance coefficient. However, empirical evidence on the evolution of structural heterogeneity in the economy since the 1960s is notoriously scarse (Head and Mayer (2013)).

We pursue the idea that a key parameter for understanding movements in the distance coefficient is the one measuring the degree of structural heterogeneity in the economy. Following Arkolakis et al. (2012) we refer to this parameter as the 'trade elasticity'. Because of data limitations inherent to our 60-year perspective, we can only estimate the trade elasticity in the Armington framework. Structural heterogeneity is determined in this framework by perceived product substitutability, i.e. the degree of product differentiation by place of origin.

We estimate the distance elasticity of aggregate trade and the Armington elasticity of substitution between goods of different origin in each year between 1962 and 2013. These two elasticities are identified separately. We deduce the implied evolution of the elasticity of trade costs to distance from the two estimated elasticities[9]. Our main result is that the increase in the Armington elasticity not only rationalizes the non-decreasing distance elasticity of trade but also hints at a reduction in the elasticity of trade costs to distance. We conclude that in the Armington framework the distance puzzle is fully explained by the increasing sensitivity of consumers to price differences. More generally, our results suggest that the distance puzzle may be due to a reduction in structural heterogeneity.

To the best of our knowledge, Berthelon and Freund (2008) is the only paper that studies the impact of changes in perceived product substitutability on the distance coefficient. Using estimates of sectoral Armington elasticities obtained by Broda and Weinstein (2006), Berthelon and Freund (2008) find a positive relationship between the variation in sectoral distance coefficients and the variation in Armington elasticities between 1985-1989 and 2001-2005[10]. Our approach is different from Berthelon and Freund (2008) because we focus on the aggregate Armington elasticity and provide direct estimates of this parameter in each year between 1962 and 2013. Our approach is different from Broda and Weinstein (2006) because we exploit cross-sectional variation in prices and trade shares to infer the extent of substitutability while these authors exploit changes in prices and trade shares over time to identify one point estimate. Our approach enables us to trace out changes in the perceived similarity of the product mix that countries supply to the world market.

We proceed in three steps. First, we examine the hypothesis that the distance puzzle is a by-product of compositional changes in the set of traded goods or of trading pairs. We refute this hypothesis by showing that the distance puzzle is more pronounced in the stable set.

[9] Erkel-Rousse and Mirza (2002) do this exercise for just one point in time on a subsample of world trade flows.

[10] Berthelon and Freund (2008) work with 776 sectors defined at the SITC Rev.2 4-digit level. Broda and Weinstein (2006) use time-series variation in prices and market shares for the set of exporters to the US market to get one value for the Armington elasticity in 1972-1988 and another value in 1990-2001.

Second, we point out that a straightforward nonlinear estimation approach allows identifying the aggregate Armington elasticity. The key intuition follows Imbs and Méjean (2015) who show that a consistent estimate of this elasticity is obtained by implementing a non-linear estimator on sectoral data while constraining sectoral elasticities to equality. In the estimation we instrument unit values - our proxy of prices - with lagged unit values together with the lagged real exchange rates specific to each bilateral relationship to address endogeneity concerns. Third, we present the results of this method and conduct a number of robustness checks.

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1 The magnitude of the distance puzzle

In this section we evaluate the sensitivity of the distance puzzle to compositional changes in the set of traded goods and in the set of trading pairs over 1962-2013. Such changes were identified in previous estimations of the loglinearized gravity model as explanatory of movements in the distance coefficient[11]. We estimate the gravity model in multiplicative form and document that the distance puzzle is magnified in the sample of stable pairs and robust to fixing the sectoral composition of world trade.

[11] Berthelon and Freund (2008) find that changes in the sectoral composition of world trade do not help to explain movements in the distance coefficient in 1985-2005. Head and Mayer (2013) find that the distance puzzle is reduced in the set of stable pairs between 1960 and 2005 when the model is estimated in loglinear form.

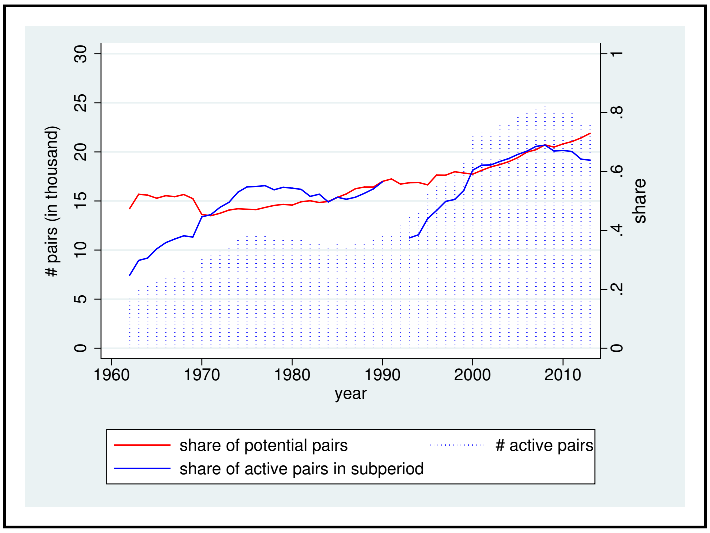
1.1 The magnitude of the sample composition effect

We use the COMTRADE dataset to make our investigation of the distance puzzle directly comparable to Head and Mayer (2013) and Berthelon and Freund (2008). We work with the 4-digit SITC Rev.1 product classification (600-700 goods) because it provides the longest and most comprehensive coverage of disaggregate bilateral trade (1962-2013). Data on bilateral distance, bilateral trade cost controls such as adjacency, common language, colonial linkages, and data on belonging or having once belonged to the same country are taken from the CEPII[12].

We conduct the estimation on CIF import flows. We restrict the sample to trade in goods which are attributed to specific 4-digit categories and to pairs for which we have data on bilateral trade cost controls. App. A lists the resulting set of countries. For each active pair attributed sectoral flows are summed to obtain total bilateral trade. We refer to the resulting sample as 'the full sample'. It covers between 88% and 99% of reported trade in COMTRADE.

Fig.1 summarizes the main features of the data. The number of active pairs increases more than fourfold in 1962-2013 (in dash, left scale), both because more countries report trade to COMTRADE and because more pairs have non-zero trade flows (Helpman et al. (2008)). Active pairs make up between 45% and 73% of the total number of possible trade relationships, with a clear upward trend (in red, right scale).

Figure 1: Active pairs in COMTRADE (1962-2013)



If we focus on the set of pairs that report non-zero trade in at least one year, the share of

[12] See Mayer and Zignago (2011). The database is available at www.cepii.fr. We constructed bilateral distance and bilateral cost controls for East and West Germany, USSR, and Czechoslovakia.

active pairs increases by 22 percentage points between 1962 and 1990 and by 27 percentage points between 1993 and 2013 (in blue, right scale)[13]. By the end of the sample about 2/3 of pairs which trade at least once between 1962 and 2013 are reporting non-zero trade.

Hence, sample composition effects are substantial. Nonetheless, the bulk of total trade is attributable to the 786 pairs which trade both ways in every year. We refer to this set of stable reciprocal pairs as the 'superbalanced' sample and use it to investigate the magnitude of the sample composition effect[14].

We follow the canonical Anderson and van Wincoop (2003) derivation of the gravity model to express aggregate bilateral trade X_{ij} as a function of bilateral trade barriers τ_{ij}, multilateral trade resistance terms in source i and destination j (resp. Π_i and P_j), and nominal incomes Y_n with $n \in \{i, j, w\}$ where w is world income[15].

$$X_{ijt} = \left(\frac{Y_{it}Y_{jt}}{Y_w t}\right)\left(\frac{\tau_{ijt}}{\Pi_{it}P_{jt}}\right)^{\epsilon_t} \tag{1}$$

We include the time subscript t not only on each variable but also on the elasticity of trade flows to trade costs ϵ to underline that this parameter is subject to change. In the Armington framework $\epsilon_t = 1 - \sigma_t$ where σ_t corresponds to the elasticity of substitution between goods of different national origin. We seek to quantify the evolution of the elasticity of aggregate trade flows to trade costs. Hence, the key parameter of interest for this paper is the Armington elasticity which captures perceived substitutability of composite goods which differ by their place of production.

As total bilateral trade costs τ_{ijt} are not directly observed for each pair and year, we model them as a function of observable time-invariant bilateral controls which are distance, adjacency, and common language together with persistent but time-varying controls standard in the gravity literature which are historical and current colonial linkages as well as belonging or having once belonged to the same country. We include an unobserved bilateral trade cost component ν_{ijt} assumed to have mean zero conditional on the observables[16]. We denote distance $\eth_{ij}$, group the other time-invariant observables in the vector Z and time-varying observables in the vector S_t

13 We split the sample in two subperiods, 1962-1990 and 1993-2013, to take into account country creation and disappearance in the early 1990s.

14 The superbalanced sample includes 32 countries listed in App. A. The appendix discusses trade coverage.

15 This formulation is not specific to the Armington framework. See footnote 20 in Eaton and Kortum (2002) and subsequent discussions of equivalence in Arkolakis et al. (2012); Head and Mayer (2013).

16 The error term contains bilateral variation in trade costs due to trade policy. The question of possible changes in the distance distribution of trade costs as a consequence of policy decisions is beyond the scope of this paper.

to get the following specification of the trade cost function:

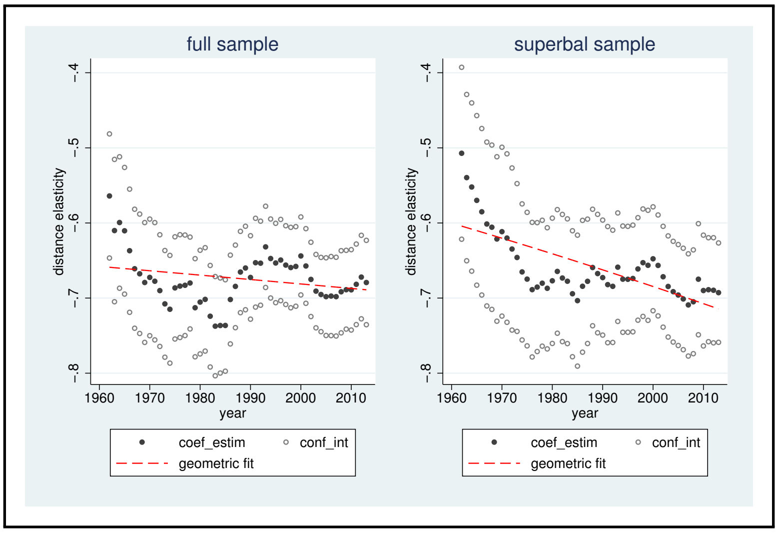
$$\tau_{ijt} \quad = \quad \exp\left\{\rho_t \ln \eth_{ij} + Z'\zeta_t + S_t{}'\varsigma_t + \nu_{ijt}\right\} \tag{2}$$

Replacing (2) in (1), substituting source and destination specific variables with country fixed effects (resp. f_{it} and f_{jt}), defining a constant ξ_t and specifying a multiplicative error term ξ_{ijt} which includes the exponentiated unobserved bilateral trade cost gives the equation to be estimated on aggregate bilateral trade:

$$X_{ijt} \quad = \quad \exp\left(\xi_t - \delta_t \ln \eth_{ij} + Z'\zeta_t + S_t{}'\varsigma_t + f_{it} + f_{jt}\right)\xi_{ijt} \tag{3}$$

To ensure consistency of the point estimates we do not loglinearize the model although switching to a non-linear estimator may entail a loss of efficiency (Manning and Mullahy (1999)). We implement (3) in the full and superbalanced samples using the PPML estimator (Silva and Tenreyro (2006)). The estimation is conducted in cross section. The parameter of interest is the distance elasticity, $-\delta_t$, which corresponds to the product of the distance elasticity of trade costs ρ_t and of the trade elasticity ϵ_t[17].

Figure 2: The sample composition effect



Results for both samples are shown in fig.2. In the full sample (left pane) the distance effect δ_t increased by 4.5% between 1962-2013 (left pane), but this increase is only marginally significant. In the superbalanced sample (right pane) the increase in the distance effect δ_t is magnified to 18.2%, and this increase is strongly significant[18]. Consistently with previous

[17] We follow the notation in Head and Mayer (2013) (see equation (4), p.1205).

[18] The annualized growth rate is 0.09% in the full sample and 0.33% in the superbalanced sample.

studies we find that the distance sensitivity of trade has not been reduced over time. Moreover, the distance puzzle is enhanced in the sample of stable trade relationships[19].

We obtain the large confidence intervals shown in fig.2 when we implement the Huber-White correction of standard errors[20]. Standard errors are reduced by an order of magnitude when this correction is not implemented. Considering that the PPML approach already takes into account heteroskedasticity, and that we are looking at these point estimates as if they were descriptive statistics on the whole population, we take the results shown in fig.2 as sufficient evidence of the robustness of the distance puzzle to sample composition effects.

Next, we demonstrate that the distance puzzle is robust to product composition effects. In particular, the non-linearity in the evolution of the distance coefficient disappears once we control for product composition effects (see fig.4 below).

1.2 The magnitude of sectoral composition effects

We assess the incidence of sectoral composition effects in two ways. The first exercise consists in fixing the sectoral composition of world trade. The second exercise consists in fixing the sectoral composition of the bundle supplied by each exporter to the world market.

In the first exercise we fix the sectoral composition of total trade to the initial year of the sample. Denoting each 4-digit sector k and the annual share of the sector in world trade (w,t) by $s^k_{w,t}$, the reweighting procedure fixes the share of each 4-digit sector in world trade to its share in 1962. The reweighted sectoral bilateral flow is $\tilde{X}^k_{ijt} = X^k_{ijt} * \frac{s^k_{w,1962}}{s^k_{w,t}}$. The reweighted sectoral flows are summed for each pair, and the gravity equation is estimated in each year for aggregate bilateral trade.

Results are shown in fig.3. The evolution of the distance coefficient becomes much more linear in the full sample (left pane), and this exacerbates the distance puzzle to a 14.5% increase in δ_t. The impact of this reweighting procedure is about nil in the sample of stable pairs (right pane) where δ_t increases by 18.4%[21]. Indeed, the main incidence of fixing the sectoral composition of world trade is the elimination of short-term fluctuations in the distance coefficient due to fluctuations in the weight of the energy sector. As this sector plays a relatively minor

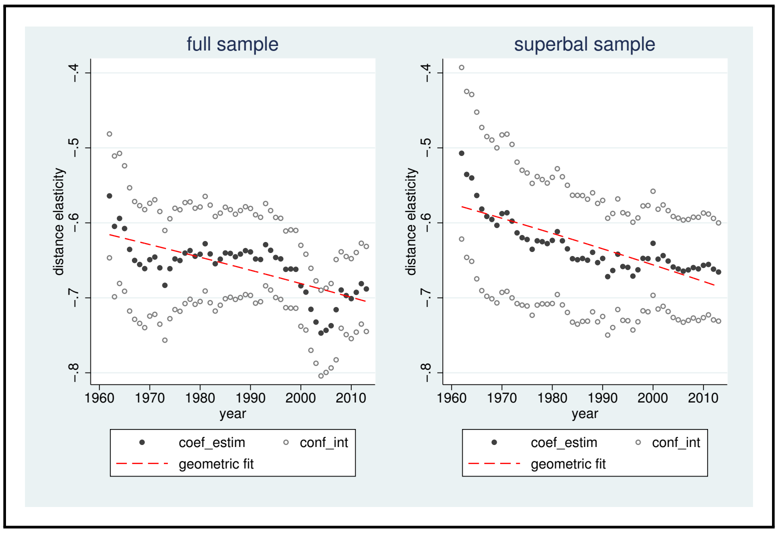
[19] The opposite result holds in the loglinear specification (Head and Mayer (2013)). In OLS the decrease in the distance elasticity due to the elimination of small trade flows between distant partners trumps the increase in the distance elasticity of stable trade relationships.

[20] Switching to the panel approach with the full set of country-year dummies does not achieve significant improvement in precision.

[21] The annualized growth rate is 0.26(0.33)% in the full (superbalanced) sample. Confidence intervals are reported for the Huber-White correction of standard errors. Without it standard errors are reduced by a factor of 10.

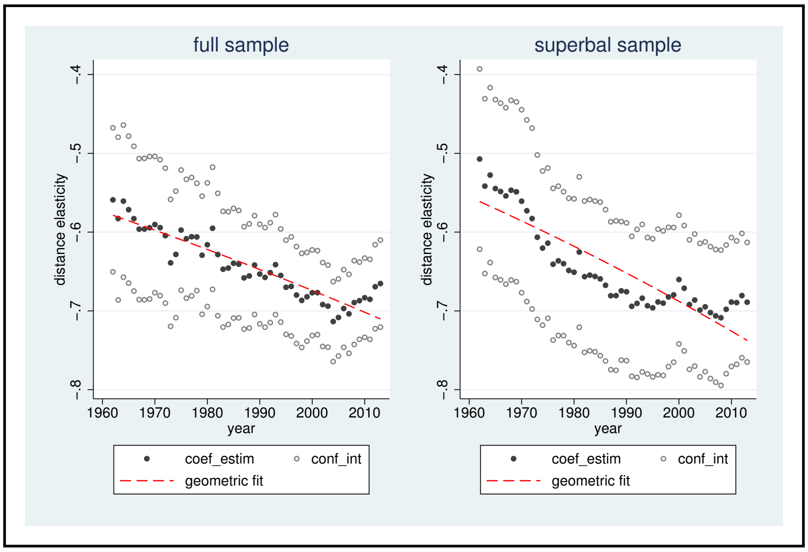
role in trade of stable reciprocal partners, the reweighting procedure has little incidence on the distribution of trade over distance in this sample.

Figure 3: Product composition effect: fixing the world bundle



In the second exercise we fix the composition of the bundle supplied by each exporter i to the world market. Denoting the annual share of the sector in world imports from i by $s^k_{i,t}$, the reweighting fixes the share of each 4-digit sector in world imports from i to its share in 1962. The reweighted sectoral bilateral flow is $\tilde{X}^k_{ijt} = X^k_{ijt} * \frac{s^k_{i,1962}}{s^k_{i,t}}$. The resulting sectoral flows are summed for each pair, and the gravity equation is estimated on total reweighted bilateral trade.

Figure 4: Product composition effect: fixing the country bundle



As illustrated in fig.4, fixing the composition of the country-specific composite good exacerbates the magnitude of the distance puzzle. Further, while the degree of precision in the estimation of the gravity equation with the Huber-White correction of standard errors remains similar to the benchmark specification, 87% of the variation in the distance coefficient is attributable to the time trend in the full sample, against just 7% in the benchmark specification[22]. The corresponding annualized growth rate is .40% in the full, and .54% in the stable sample. This corresponds to a 22.7% increase in the distance sensitivity of trade between 1962 and 2013 in the full sample, and to a 31.4% increase in the sample of stable pairs.

We conclude that short-term fluctuations in the distance effect δ_t are likely attributable to product and sample composition effects. But the long-term evolution of the distance effect appears linked to structural changes. These structural changes are more pronounced in the set of stable trade relationships.

1.3 Summing up: the robustness of the distance puzzle

Table 1: Evolution of δ_t: sample, composition and FTA effects

	FULL			STABLE		
	rate (%)	R-sq	total change	rate (%)	R-sq	total change
Baseline	.09*	.07	1.045	.33***	.49	1.182
World bundle	.26***	.53	1.145	.33***	.68	1.184
Country bundle	.40***	.87	1.227	.54***	.77	1.314

Note: Estimated annualized growth rates reported in col.2 and col.5 are obtained as a geometric fit on the basis of annual point estimates of the distance coefficient in 1962-2013. Col.3 and col.6 report the share of time variation in the point estimate explained with the annualized growth rate.

Table 1 summarizes our findings. The distance puzzle is magnified in the sample of stable pairs and robust to fixing the product composition of world trade. Hence, the non-decreasing distance elasticity of trade is likely to be a structural outcome rather than an artefact of composition effects. These results motivate our focus on structural heterogeneity as a possible alternative explanation of the non-decreasing distance elasticity.

[22] In the superbalanced sample 77% of the variation is attributable to the time trend when the product bundle is fixed against 49% when the product bundle is not fixed.

2 Interpreting the distance coefficient

2.1 What does the trade elasticity actually measure?

The distance coefficient is the product of two elasticities: the elasticity of trade costs to distance rho and the elasticity of trade flows to trade costs $zeta$. The question addressed in this paper is very simple. Have trade flows become more sensitive to trade costs (increasing $zeta$), or have trade costs become more sensitive to distance (increasing ρ)? In section 3 we estimate the evolution of $ zeta $ to deduce the evolution of $ rho $. The three main microfoundations of the gravity model of trade give structurally different interpretations to $ zeta $ but not to $ rho $. In this section, we provide details on the procedure used in this paper to estimate the evolution of $ zeta $ in the Armington framework.

In (Eaton et Kortum 2002) the heterogeneity dimension captured by the trade elasticity $ zeta $ is intersectoral. Consumers maximize a CES utility function defined at the intersectoral level by shopping around the world for the cheapest supplier of each sectoral good.

In the Melitz-Chaney framework the heterogeneity dimension captured by the trade elasticity $ zeta $ is intrasectoral. Consumers value firm-specific varieties of sectoral goods which they acquire in monopolistic competition markets.

In (Anderson et van Wincoop 2003) there is no heterogeneity in productive efficiency. The production process in each country and sector is constant returns to scale. There is thus perfect competition between domestic producers. The heterogeneity dimension comes from the assumption that consumers perceive products of different national origin as intrinsically imperfect substitutes. Each country is specialized in production of country-specific varieties which on aggregate give a country-specific composite good. The parameter $ zeta $ is the lower tier Armington elasticity of substitution. It measures the degree of substitutability of goods of different national origin.[23]

2.2 What do we know about the evolution of the trade elasticity?

In neither of these models there is a theoretical mechanism to explain a change in the trade elasticity overtime. A shock to consumer preferences or to the shape parameter of the productivity distribution would be required. Nonetheless, the heterogeneity parameter measured

[23] The upper-tier elasticity measures substitutability of domestic products and an aggregate import good. The lower-tier one measures substitutability between importers of a given good. See (Sato 1967; Reinert et Shiells 1991; Saito 2004).

on aggregate trade data could have evolved over 1962-2009 without any shock to the underlying heterogeneity, either through changes in the range of traded goods, time-sensitive aggregation issues linked to estimating a single parameter across sectors, or agents' adaptation to a changing economic environment. However, there is little empirical evidence on the evolution of sector-specific and aggregate trade elasticites in either model.

To the best of our knowledge, there is no direct empirical evidence on the evolution of sectoral or aggregate efficiency dispersion parameters in the heterogenous firms framework. Different theoretically grounded methods have been used to estimate aggregate trade elasticities in the Ricardian framework for a specific year ((Eaton et Kortum 2002; Simonovska et Waugh 2014; Costinot, Donaldson, et Komunjer 2012; Caliendo et Parro 2012), but only (Levchenko et Zhang 2011) study the evolution of intersectoral productivity dispersion. They find evidence of within-country convergence in sectoral knowledge stocks in 1960-2010. As there is less heterogeneity in producer efficiency across the set of goods comparative advantage exerts a weaker force against trade resistance imposed by trade barriers.

Evidence on the evolution of lower-tier Armington elasticities of substitution measured on aggregate data is scarce. For France, (Welsch 2006) provides estimates of aggregate lower-tier Armington elasticities since the 1970s. He finds that among exporters to the French market the elasticity peaked in the 1970s, and progressively decreased in the 1990s. (Broda and Weinstein (2006)) provide evidence on the evolution of sectoral Armington lower-tier elasticities between 1972-1988 and 1990-2001 for American imports. They find that they have decreased for all types of goods at all levels of product disaggregation, i.e. at the 10-digit, 5-digit, and 3-digit levels. These results indicate that the parameter estimated on aggregate trade data would also have decreased, deepening the distance puzzle. But to the best of our knowledge, no paper has as yet provided evidence on the evolution of Armington elasticities for aggregate bilateral trade while constraining the parameter to be the same across destination markets.

2.3 A method to measure trade elasticity in the Armington framework

Measuring Armington elasticity is an perennial issue in trade litterature. Feenstra et al. (2014) discusses the difficulties and a state-of-the art method to measure both higher-tier (or "macro") Armington elasticities and lower-tier (or "micro") ones (see also Feenstra (1994), and refinements in Broda and Weinstein (2006) and Imbs and Méjean (2013)). We depart from this litterature and suggest a cruder estimation method for numerous reasons. First, as we want to measure the lower-tier Armington elasticity from 1963 for all the countries involved in world trade, we operate in a data-poor environement. We cannot use dissagregated domestic prices nor production. We believe that the biases entailed by this lack of information do not make the exercice worthless because we are interested in the effect of the evolution of the parameter rather than in its exact value. Second, Feenstra's method and its developments relies on the assumption the elasticicy parameter remains constant through time, whereas our research question implies that the parameter can vary from year to year. Third, more fundamentally, Feenstra's elasticity parameter determines short-run, marginal, longitudinal effects whereas we are interested in the elasticity parameter which determines long-run, equilibrium, cross-section outcomes. While we admit that in most tracktable theoretical settings these would be the same. In the absrtact, is not immediate that they should be.

The UN COMTRADE bilateral trade database covers the majority of countries over 1962-2009. It gives information on trade flows and cif unit values at the SITC 4-digit level. This data are sufficient to estimate the trade elasticity in the Armington framework. If we have importer-specific prices in destination markets and importer-specific market share, we should be able to observe some statistical regularities. The basic intuition of the method we use starts from the well-known result that assuming CES utility function in the one-good Armington framework we have:

$$X_{ij} = \left(\frac{P_{ij}}{P_j}\right)^{-(\sigma-1)} Y_j$$

where X_{ij} is the cif value of the exports from i to j, P_{ij} is the cif price of the good shipped from i to j and P_j is the price index in the destination and Y_j total import demand in the destination. The exponent $(\sigma - 1)$ captures substitutability of country-composite goods across frameworks. It is also the aggregate trade elasticity ζ in the Armington framework. Briging this equation to the data is difficult, however, as we do not observe aggregate prices, but unit values at the SITC 4-digit category level. Still, the distance puzzle concerns an elasticity estimated on aggregate trade data. As shown by (Imbs et Méjean 2013) this parameter cannot generally be mimicked by a theoretically grounded weighted average of sector-specific trade elasticities. Hence, we need an estimation procedure that works directly with aggregate data. Define aggregate imports from source country i to destination country j as the sum of imports from each sector k where a sector corresponds to a SITC 4-digit category: $X_{ij} = \sum_k X_{k,ij}$. Given CES utility at the intersectoral level, sectoral demand in country in sector for imported goods is given by:

$$Y_{k,j} \quad = \quad \left(\frac{P_{k,j}}{\beta_k P_j}\right)^{1-\sigma} Y_j$$

Where $P_{k,j}$ and P_j are price indexes, $\beta_k > 0$ is a sector-specific preference parameter, Y_j is total demand for imported goods, $\sigma > 1$ is the elasticity of substitution between sectors. Assume each country exports a specific national variety. Preferences within each sector k between national varieties are assumed well represented by a CES utility function with the same σ parameter as the intersectoral CES utility function. Intrasectoral demand for varieties exported by i in j in sector k is:

$$X_{k,ij} = \left(\frac{p_{k,ij}}{\gamma_i P_{k,j}}\right)^{1-\sigma} Y_{k,j}$$

Where $\gamma_i > 0$ is a origin-country-specific preference parameter and $P_{k,j}$ is the CES price index:

$$P_{k,j} = \left[\sum_{i \neq j} \left(\frac{p_{k,ij}}{\gamma_i}\right)^{1-\sigma}\right]^{1/(1-\sigma)}$$

Defining $\frac{Y_{k,j}}{Y_j} = \omega_{k,j}$, we get:

$$\frac{X_{k,ij}}{Y_j} = \omega_{k,j}\left(\frac{p_{k,ij}}{\gamma_i P_{k,j}}\right)^{1-\sigma}$$

Summing over all SITC 4-digit sectors:

$$\sum_{k=1}^{K} \frac{X_{k,ij}}{Y_j} = \frac{X_{ij}}{Y_j} = \gamma_i^{\sigma-1} \sum_{k=1}^{K} \omega_{k,j}\left[\frac{p_{k,ij}}{P_{k,j}}\right]^{1-\sigma}$$

The market share equation for aggregate bilateral trade as a function of the weighted average of sectoral relative prices of in is:

Changing notations:

$$\frac{X_{ij}}{Y_j} = \kappa_i \sum_{k=1}^{K} \omega_{k,j} \frac{p_{k,ij}^{1-\sigma}}{\sum_{l \neq j} \kappa_l p_{k,lj}^{1-\sigma}} \tag{4}$$

$$\tag{5}$$

As this is a market share, it is reasonable to assume that the errors are multiplicative. Measurement errors on a market share of 1% cannot be the same in percentage points as measurement errors on a market share of 10%. We have to fit the following model on the data:

$$\frac{X_{ij}}{Y_j} = \kappa_i \sum_{k=1}^{K} \omega_{k,j} \frac{p_{k,ij}^{1-\sigma}}{\sum_{l \neq j} \kappa_l p_{k,lj}^{1-\sigma}} .e^{\varepsilon_{i,j}} \tag{6}$$

Notice that we only have non-zero observations, as $p_{k,ij}$ is only observed when there is a trade flow. We take logs to transform the errors into additive ones and estimate the following equation for each year with a non-linear least square procedure in STATA:

$$\ln\left(\frac{X_{ij}}{Y_j}\right) = \ln(\kappa_i) + \ln\left(\sum_{k=1}^{K} \frac{Y_{k,j}}{Y_j} . \frac{p_{k,ij}^{1-\sigma}}{\sum_{l \neq j} \kappa_l p_{k,lj}^{1-\sigma}}\right) + \varepsilon_{i,j} \tag{7}$$

This yields annual estimates of κ_i and σ[24].

3 Evolution of the Armington trade elasticity in 1963-2009

3.1 The incidence of missing unit values

To estimate the market share equation on the COMTRADE dataset we need to tackle the question of missing information on trade flows and unit values (uv).

A first difficulty arises when the trade flow is observed but information on quantities is missing, and it is therefore not possible to compute the unit value. Figure 5, *4'-digit* indicates how often that happens from 1962 to 2013. The problem arises for a small share of the total trade value (never more than 25%). We assume that information on quantities is missing due to imperfections in the data collection procedure, and that bilateral trade flows are observed with a similar degree of precision whether or not quantity has been recorded.

To deal with missing uv, we impute prices from similar products using a stepwise price imputation procedure .

The stepwise price imputation procedure is as follows. The relative price of each source in the destination is constructed at the highest disaggregation level for each product and quantity unit in which the source is active, the *4'-digit* level. We then proceed level by level for aggregation: the relative price of the composite sectoral good of the source is constructed at the 4-digit level using the weighted average relative price observed at the 4'-digit level, with destination-specific weights for each variety of the 4'-digit good the source is active in. If no price is available at the 4-digit level, we move at the 3-digit level, etc. As Figure 5 shows, that procedure allows to reduce considerably the share of trade without unit values. This improves the estimation of prices if one assumes that missing destination-specific relative prices at the 4'-digit can be approximated by the mean observed destination-specific relative price among the corresponding 4-digit group (and similarly at each aggregation level).

3.2 Zero trade flows

A second difficulty arises when both quantity and trade data are missing. Zero trade flows (ztf) are a prevalent feature of the data even though under model assumptions some trade should

[24] Because of the way the procedure nl works in STATA, instead of minimizing $\sum_{i,j} \left(\varepsilon_{ij}\right)^2$, it minimizes $\sum_{i,j} n_{ij} \left(\varepsilon_{ij}\right)^2$ where n_{ij} is the number of 4-digit sectors in which imports from i to j are recorded. As a result, we have to weight the regression by $1/n_{ij}$ to make its results comparable to those in part one.

Figure 5: Trade with missing unit values

The way to read this graph is : In 1962, 10% of trade is in a 3-digit bilateral trade category that does not mention at least one unit value be observed in every sector k between all pairs ij.[25] To resolve this contradiction, we assume that this information is missing because the underlying trade flow is positive but so small that it does not pass the threshold applied by the data collecting authorities (in UN COMTRADE this threshold corresponds to 1000 USD). Such flows, if recorded, would not substantially modify the distribution of observed market shares in the destination (the left hand side of equation 7) because they are an order of magnitude smaller than observed trade. Still, we have to make some assumption about their price to be able the effect of price on market shares.

This is only possible when zero trade flows are present in an existing bilateral relation. In this case, we use the same stepwise price imputation for zero trade flows as in the case of missing unit values. This is problematic because statistically unobserved trade values logically correspond to a higher cif price than the maximum observed price in the destination across all sources and sectors while by construction we postulate that unobserved relative prices in ztf sectors are equal to a weighted average relative price across sectors in which bilateral trade is observed. [26]

This assumption would not bias our estimate if the underestimation factor were constant across exporters. This scalar would cancel out across sources, and the estimated substitutability parameter would correspond to the true parameter. Table 2 shows it is not the case. The share of ztf is strongly decreasing in market share, i.e. the underestimation factor is larger for small exporters (though they already have higher prices). As a result, for a given observed distribution of market shares, the underlying dispersion in relative prices of the composite good is greater than the observed dispersion in relative prices. This means that the estimated parameter $(1-\tilde{\sigma})$ overestimates the true substitutability parameter $(1-\sigma)$.

Table 2 shows that the reduction in the share of ztf proceeds at quicker pace in 1962-2009 for small exporters: the coefficient for the interaction term for the market share and year is significant and positive. Table 3 presents the predicted share of ztf for four types of exporters in 1962 and 2009. For a very small exporter with .02% market share, the initial share of ztf is predicted to be .95, and it is reduced to .83 by 2009, i.e. a 12 percentage point decrease. Consider a relatively big exporter, with a 10% market share: its share of ztf is reduced from .72 to .65, a 7 percentage point decrease. As the gap between the share of ztf for big and small exporters is reduced overtime, the overestimation bias of $(1-\tilde{\sigma})$ is progressively reduced.

[25] The share of observed SITC 4-digit flows relatively to the total number of potential SITC 4-digit bilateral trade flows increases from 10% to 14% in 1962-2009.

[26] An alternative method consists in imputing unobserved relative prices with some arbitrary price above the maximum observed in the destination. As ztf constitute 85-90% of all 4-digit trade flows, this method is problematic because results are driven by imputed rather than observed prices.

Table 2: Proportion of zero trade flows as a function of market share

Proportion of zeros as a function of market share

VARIABLES	(1) propor_ssuv_5	(2) propor_ssuv_5	(3) propor_ssuv_5	(4) propor_ssuv_5
ln_ms	-0.0330***	-0.1261***	-0.0350***	-0.1542***
	(0.0001)	(0.0098)	(0.0001)	(0.010)
year	-0.0026***	-0.0022***	-0.0030***	-0.0025***
	(0.0000)	(0.0001)	(0.0000)	(0.000)
interaction		0.0000***		0.0001***
		(0.0000)		(0.000)
Constant	4.8436***	4.0417***	5.6383***	4.6135***
	(0.0244)	(0.1007)	(0.0270)	(0.098)
Observations	749,686	749,686	749,686	749,686

Robust standard errors in parentheses

*** p<0.01, ** p<0.05, * p<0.1

The proportion of zeros is computed at the SITC 5-digit level.

CETTE TABLE EST À VÉRIFIER / REPRODUIRE POUR LA BONNE PÉRIODE.

Notes: The share of ZTF is computed at the SITC 4-digit level. The estimation is conducted in PPML in order to include observations where ztf=0. The log of the market share is used in the estimation. Destination fixed effects are included in (3) and (4). Robust standard errors are in parentheses. *** p<0.01.

Table 3: Predicted share of ztf for exporters with different market share, 4-digit level

Table 2: Predicted share of ztf for exporters with different market share, 4'-digit level

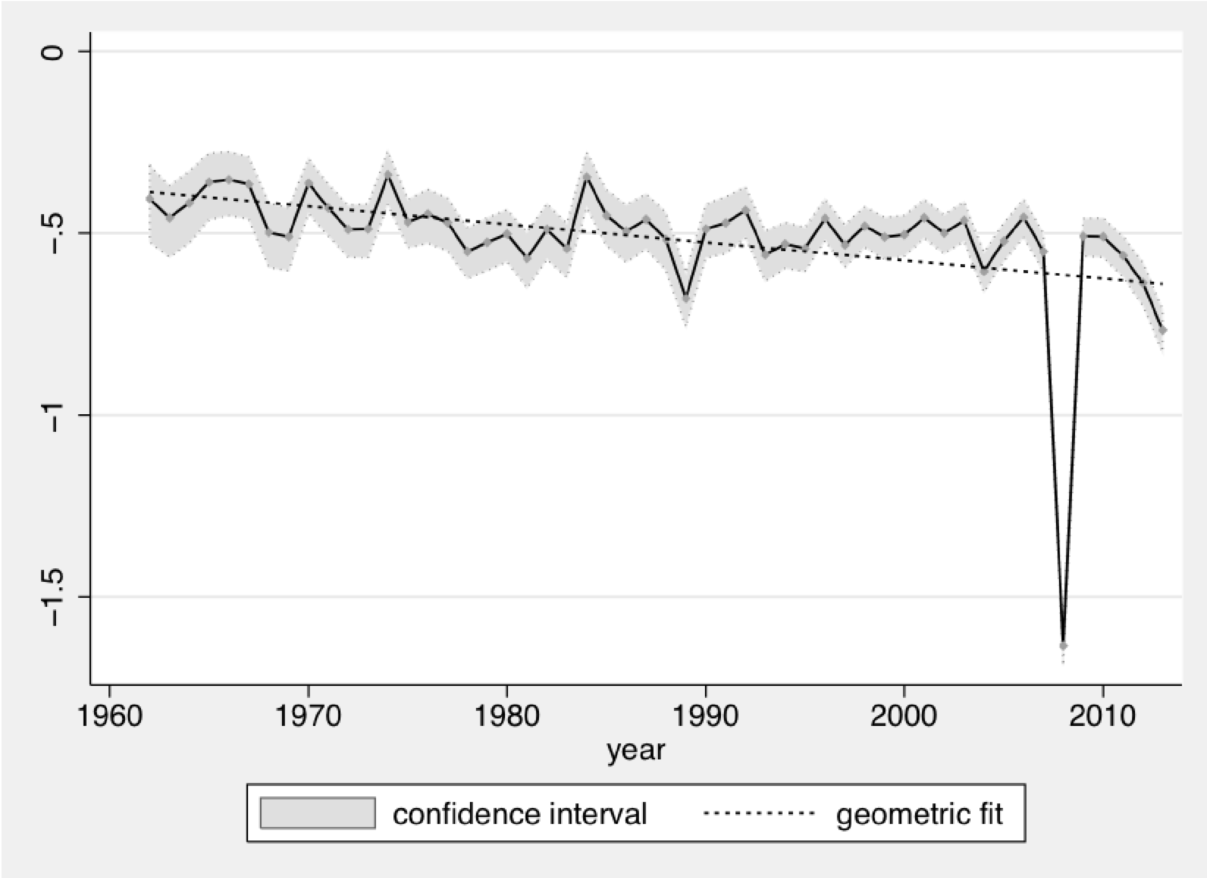
year	expmean	exponepct	exptenpct	exptwostdv
1962	0.96	0.84	0.78	0.75
2009	0.85	0.75	0.70	0.67

Columns (1) and (4) correspond to the mean and to 2 st. deviations above the mean in the distribution of log market share. Columns (2) and (3) correspond to the mean and to 2 st. deviations above the mean in the distribution of market share.

Thus, the hypothesis we make on unobserved sectoral prices in ztf sectors does not always impede interpreting the evolution of the underlying substitutability parameter. In particular, because the overestimation bias is reduced overtime, if it is found that the estimated parameter increases in absolute value, this evolution necessarily provides a lower bound on the increase in the underlying substitutability parameter.

Figure 7 presents the results on the evolution of $(1-\tilde{\sigma})$ obtained when equation 7 is estimated in annual cross-sections of the COMTRADE dataset. 2008 is obviously an outlier. However, even excluding 2008, the absolute value of trade elasticity has increased by 38% from 1962 to 2013. This corresponds to an annual increase of .6% per year.[27] According the preceding discussion, this is a lower bound on the increase in the underlying substituability parameter.

Figure 7: Estimated $(1-\tilde{\sigma})$



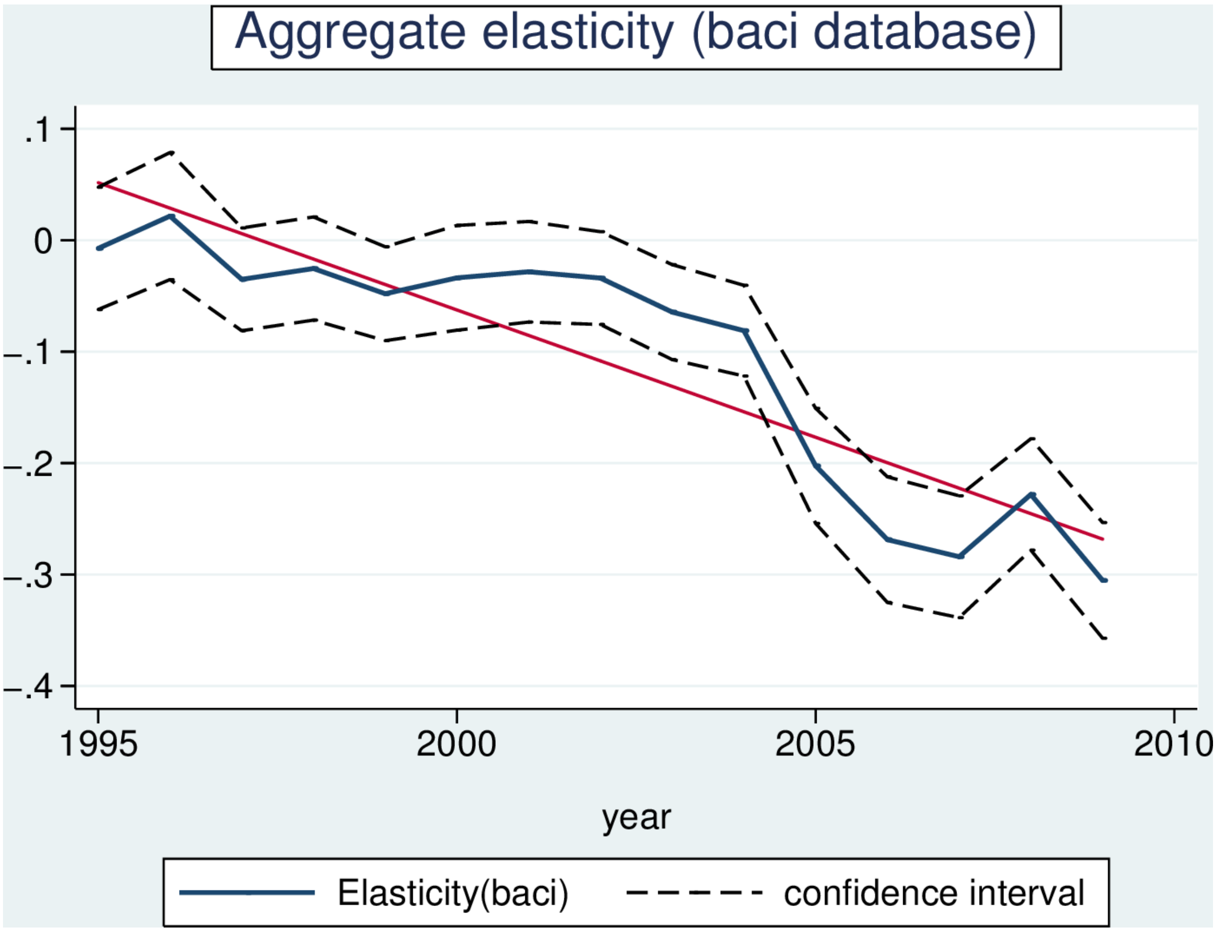
3.3 Robustness checks

3.3.1 Changing the dataset

We provide a robustness check by estimating the evolution of the heterogeneity parameter for aggregate bilateral trade on a different dataset. We use the BACI dataset which reports bilateral trade data at the HS-1992 6-digit disaggregation level for 1995-2009. The accuracy of the relative prices of country-composite goods constructed with this dataset is improved because the harmonization procedure applied by (Gaulier et Zignago 2010) in constructing BACI yields much better-quality unit values while substantially reducing the number of observations with lacking unit value. As a result, at the 6-digit level, less than 7% of total reported trade in BACI

[27] The coefficient of the geometric fit is significant at 1% level, the 95% confidence interval is between 0.4% and 0.9% (not taking into account though the uncertainty around yearly estimates).

Figure 6: Figure 8 shows that our results hold: the elasticity parameter is found to increase in absolute value from 1995-2009. This can be compared with the equivalent period in our original dataset: the increase in the elasticity is much steeper on the BACI dataset. This finding supports the idea that our benchmark estimation likely provides a lower bound on the increase in the aggregate trade elasticity. However, the level of the elasticity estimated in 1995-1999 on BACI data is puzzling and suggests the existence of an attenuation bias. This is the focus of our second robustness check.



has missing unit values. This is reduced to 1-3% of total trade when the data is aggregated to the 4-digit level, as opposed to more than 10% in the raw COMTRADE data we originally used. Another advantage is that the share of ztf in BACI is stable in 1995-2009 as opposed to relatively strong fluctuations in the share of ztf overtime in our original dataset. The disadvantage of BACI is that it covers only a relatively short period compared to the years over which the distance puzzle exists. Obviously, we do not expect to reproduce exactly the results obtained with our original dataset because the trade classification and its level of aggregation are different

Figure 8: Estimated $(1-\tilde{\sigma})$ **, BACI database**

3.3.2 Price instrumentingand the non-linear estimation of σ

Instrumenting procedure

In equation , it is possible that the estimate of $sigma $ is downward-biased because of unobserved demand shifters that may be correlated with prices (Hummels et Schaur 2013). Specifically, whenever the supply curve has a finite slope, unobserved demand shocks will result in a simultaneous increase in the price and in the expenditure on sector *k* products. Further, even if the supply curve is perfectly elastic, measurement error may lead to a downward bias in the estimate of $ sigma $. This downward bias occurs whenever there is a systematic link between unobserved quality and observed unit values whereby higher observed unit values in sector *k* are associated with higher underlying quality and, consequently, higher expenditure on sector *k* products.[28]

To check whether unobserved demand shifters correlated with prices are a source of concern in the estimation, we need an instrument that adequately captures exporter-specific shocks to sector *k* prices that are not demand-driven, such as exogenous shocks to the cost of inputs. We use adjusted past unit values as an instrument for current unit value. The adjustment is based on exporter-specific price shocks as measured by domestic price evolutions. Unfortunately, we do not have information on producer price indices at the SITC 4-digit level. Even at the aggregate level, PPI information is not available for most countries and years in our sample. We therefore settle for the price level of GDP and the price level of investment.

The price level of GDP will be a valid instrument if demand shocks in the importing country are the main source of price endogeneity, and the small country assumption holds. In this case, demand shocks in any market for products of a particular exporter will have no incidence on world prices of these products and, consequently, the price level of GDP in the exporting country will be independent of idiosyncratic demand shocks in the importing country. If however the small country assumption does not hold for certain importers, the price level of GDP could itself be a function of demand for exporter-specific products. Therefore, as a robustness check, we use the price level of investment as an alternative proxy for cost shocks to production because the price of investment is more likely to be determined by global demand for industrial goods and to be exogenous to unobserved demand shocks in any given market.

[28] Another potential problem is linked to using per kg instead of per item prices. As discussed in (Hummels et Schaur 2013), unit values reflect product bulkiness, and bulkier products are more likely to be shipped via cheaper means of transportation. If products differ in bulkiness within the sector, unit value differences will overstate price differences and lead to a downward bias in the estimated elasticity.

Information on the price level of GDP and the price level of investment is reported in the Penn World Table 9.0 for 182 countries in 1950-2015 in current US dollars (R. C. Feenstra, Inklaar, et Timmer 2015). For each variable, the price level is normalized to 1 in the USA in 2005. Information is provided for 113 countries in each year, and for the remaining countries between 10 and 51 years are covered. Thus, our sample of countries does not coincide perfectly with the countries in the Penn World Table. However, as it is the smallest exporters that drop out, the sample adjustment in terms of world trade coverage is minor.

The instrumenting procedure consists in exploiting past information on unit values and on changes in the cost of domestic inputs to replace observed unit values $p_{k,ij,t}$ with predicted unit values $\hat{p}_{k,ij,t}$. We assume a stylized cost function in which the sector-specific cost is determined by an economy-wide cost measure C_{it} together with sector-producer specific characteristics summarized by the index $z_{k,i,t}$ (> 0) and bilateral characteristics of trade costs $z_{ij,t}$ (> 0). Denoting the time lag by l, we assume:

$$c_{k,ij,t} = C_{i,t}.z_{k,i,t}.z_{ij,t}$$

We assume further that the cost component is well captured by the GDP or investment price level: $C_{i,t} = P^{v}_{i,t}$ where $v = \{gdp, i\}$. Prices are a more or less sensitive to costs, depending on the competition conditions in each sector. The elasticity of prices to costs is

$$\alpha_{k,t}$$

(>0). It is a measure of the path-through that depends on specific sector conditions. Prices are also dependent on consumer-sector-producer specific characteristics $z_{k,ij,t}$ (> 0) that capture, for example, shocks to demand in j for products from i in sector k. As a result:

$$p_{k,ij,t} = \left(P^{v}_{i,t}.z_{k,i,t}.z_{ij,t}\right)^{\alpha_{k,t}}.z_{k,ij,t}$$

Then, for each time-lag l, and assuming sector-specific price-setting conditions (i.e. $\alpha_{k,t}$) constant between $t - l$ and t, we have:

$$p_{k,ij,t} = p_{k,ij,t-l}\left(\frac{P^v_{i,t-l}}{P^v_{i,t-l}}.\frac{z_{k,i,t}}{z_{k,i,t-l}}.\frac{z_{ij,t}}{z_{ij,t-l}}\right)^{\alpha_{k,t}}.\frac{z_{k,ij,t}}{z_{k,ij,t}} \tag{8}$$

$$\Leftrightarrow \tag{9}$$

$$\ln\left(p_{k,ij,t}\right) = \ln\left(p_{k,ij,t-l}\right) + \alpha_{k,t}.\ln\left(\frac{P^v_{i,t}}{P^v_{i,t-l}}\right) + \alpha_{k,t}.\ln\left(\frac{z_{k,i,t}}{z_{k,i,t-l}}\right) + \alpha_{k,t}.\ln\left(\frac{z_{ij,t}}{z_{ij,t-l}}\right) + \ln\left(\frac{z_{k,ij,t}}{z_{k,ij,t}}\right)$$

(10)

(11)

Predicted unit values are obtained by regressing observed unit values on the product of lagged unit values and the change in the price level of domestic output (investment), assuming that the sum of the three other members of the sum follows a normal law. Our baseline equation for the first stage of the estimation for a given lag is:

$$\ln p_{k,ij,t} - \ln p_{k,ij,t-l} = \alpha_{k,t} \ln\left(\frac{P^v_{i,t}}{P^v_{i,t-l}}\right) + \varepsilon_{k,ij,t,l}$$

We estimate this equation separately for each year of the sample. Hence, we allow the estimated coefficients to be year-specific. The identification assumption is that current shocks specific to the sector and to the pair ($\varepsilon_{k,ij,t-l}$) that may simultaneously drive up price and expenditure in year t are independent of lagged prices and changes in producer cost component.

The choice of the lag is associated with a trade-off. The shorter the lag, the more data we can use. However, if demand shocks in the destination are persistent, the covariance between the lagged price and the demand shock in the error term may remain positive if the first lag is used. Hence, we use a three-year lag in the baseline specifications.

For almost all years, and all specifications (there are a few exceptions at the beginning of the period) the path-through estimated in equation is statistically significant and positive. The R^2 is very small. That suggests that economy-wide domestic conditions explain a very small share of price and sector-specific price variations. Considering the dimension of the data, this is perhaps not surprising. Still, this may worry us as to the weakness of our instrument. Recall however that the instrument is actually the adjusted past unit value. Past unit values have a very strong correlation with present unit values. The correlation between adjusted past unit values and current unit values is slightly higher.

As a falsification test, we verify that variations of the unit price are not explained by past or future variations of domestic costs. We thus estimate the following equation:

$$\ln p_{k,ij,t} - \ln p_{k,ij,t-l} = \alpha_{k,t} \ln\left(\frac{P^v_{i,t}}{P^v_{i,t-l}}\right) + \beta_{k,t} \ln\left(\frac{P^v_{i,t+2}}{P^v_{i,t+1}}\right) + \gamma \ln\left(\frac{P^v_{i,t-3}}{P^v_{i,t-4}}\right) + \varepsilon_{k,ij,t,l}$$

3.3.3 Instrumenting: motivation and results (old)

The results just presented are subject to caution if supply schedules are not horizontal.[29] The demand elasticity parameter estimated in the market share equation would then be subject to attenuation bias due to not controlling for potentially positive and finite supply elasticities. This attenuation bias would not be problematic for analyzing the evolution of the substitutability parameter if only the level of the parameter were affected. The problem arises because, as shown by (R. **?**), the attenuation bias also impacts the evolution of the parameter.

As explained in 2.3, our approach is different from the canonical one as we are keen on preserving the time dimension that is central to our analysis. We need an instrument that adequately captures exporter-specific shocks to the price of the composite good which are not demand-driven, such as exogenous shocks to inputs' prices. We would like to use changes in the bilateral-specific real exchange rate. One possibility would be to use Producer Price Index (PPI) since it captures the evolution of prices faced by producers on the inputs' side. Unfortunately, we do not have PPI data for most countries and years in our sample. We therefore settle for an alternative exporter-specific price level indicator: the GDP price level in current US dollars as reported in the Penn World Tables for 189 countries in 1950-2009.[30]

The instrumenting procedure is the following. First, we compute relative prices for exporter-specific composite goods in each destination market using the stepwise price imputation procedure (see 3.1). Second, for each destination market, we compute the mean evolution of GDP price levels in current US dollars of its trading partners, weighted by their market shares in this destination. This amounts to computing the evolution of the relevant real exchange rate for each specific bilateral trade relation. Third, we compute a hypothetical relative price at time t for each exporter in each market as the product of its relative price at time $(t-1)$ and the evolution of its GDP price level between t and (t-1) relatively to all other trading partners in this destination. Fourth, we predict the relative price of each exporter in each destination at time t by regressing

[29] (Broda and Weinstein (2006)) find that supply elasticities are finite at the 4-digit level. On the other hand, (Magee and Magee (2008)) find that the small country assumption may hold in the data in which case there would be no attenuation bias.

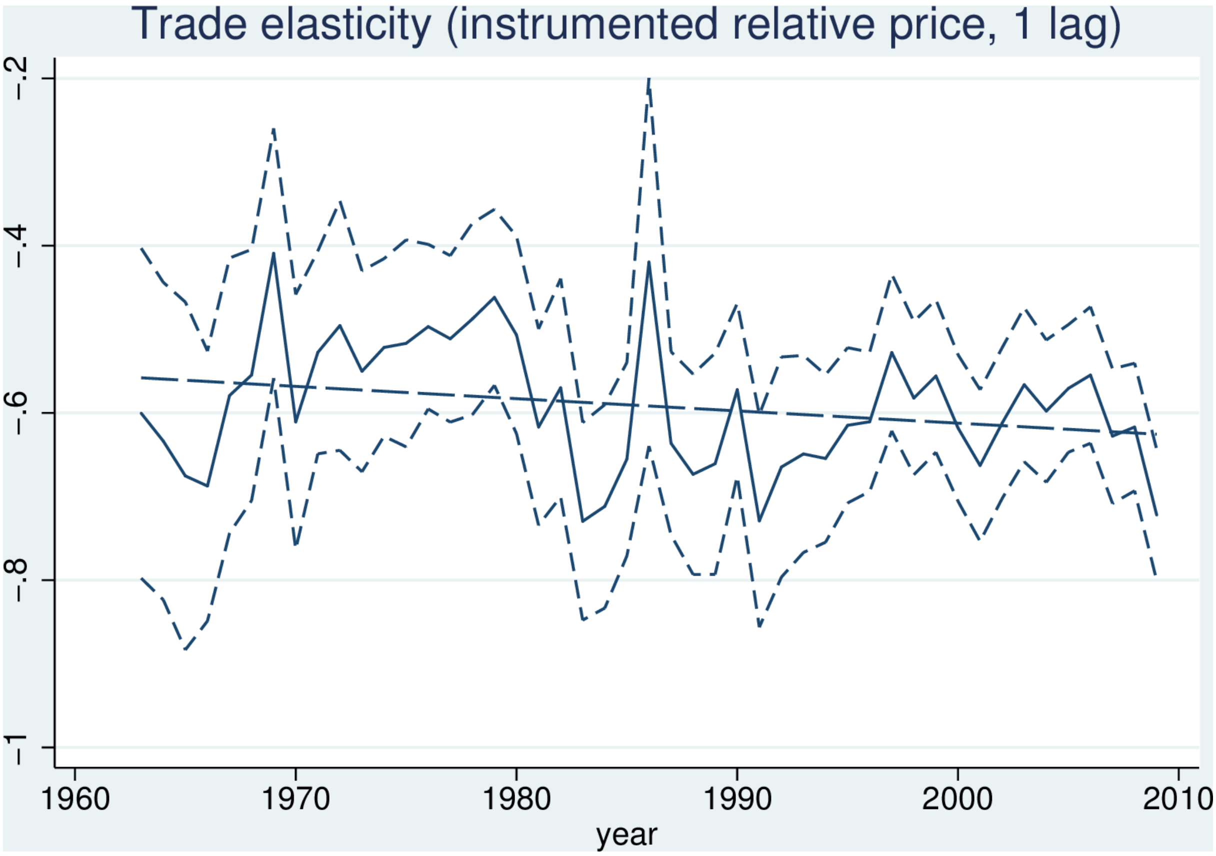
[30] See (**?**). Our sample of countries does not coincide perfectly with the countries in the Penn World Tables. However, as it is the smallest exporters that drop out, the sample adjustment in terms of world trade coverage is minor.

its observed relative price on this hypothetical relative price. This gives an instrumented relative price for each exporter which depends only on its past relative price and the relative evolution of its GDP price level. Finally, we estimate equation using these instrumented relative prices instead of the observed relative prices.

It could be argued that allowing for just one lag inadequately captures the temporal relationship between shocks to inputs' prices and their pass-through to the price of exported output. Indeed, if prices are relatively persistent, the instrumenting procedure would amount to little more than replacing observed prices in t with lagged observed prices in (t-1). We therefore also estimate equation using as instrument the evolution of each exporter's GDP price level relatively to all other trading partners in the destination between (t-s) and t where s=1,...,10.

Results obtained with one lag (s=1) are shown in Figure 9. The absolute value of the substitutability parameter has increased by 13% in 1963-2009 while the level of the estimated parameter increases by 9% on average relatively to the estimate obtained with non-instrumented prices.

Figure 9 : Estimated (1−$(1-\tilde{\sigma})$), instrumented relative price of composite good, 1 lag



This result is robust to increasing the number of periods in which the evolution of exports' prices is predicted with the evolution of domestic prices. Thus, in 1972-2009, the elasticity increases by 20% when the instrument is constructed with one lag, and by 23% when the number of lags is 10 (see Appendix **Erreur ! Source du renvoi introuvable.**). The evolution of the parameter becomes steeper as we increase the number of lags. Therefore, it is likely that our estimate provides a lower bound on the increase in the true substitutability parameter.

3.4 Is there a distance puzzle left?

This section has provided empirical evidence on the evolution of the aggregate substitutability parameter for world trade in 1963-2009. This substitutability parameter corresponds to the aggregate trade elasticity in the Armington framework. We find that this parameter has increased by 33% between 1963 and 2009 in the benchmark estimation, and by 13% when prices are instrumented. Both estimates are likely to be lower bounds on the increase in the true substitutability parameter. Section 1 has shown that the distance elasticity of trade has increased by 7% over the same period. Combining these two results, there is no distance puzzle in the framework of the Armington model in as much as the elasticity of trade costs to distance has decreased by at least 5-7% in 1963-2009.[31] Increasing perceived substitutability of country-specific composite goods contributes to the increasing distance elasticity of trade.

The reduction in the elasticity of trade costs to distance is even more pronounced if we focus on 1970-2009. As shown by (Hummels 2007), this period is characterized by a new phenomenon: the fact that air transportation starts playing a substantial role in world trade. The instrumented Armington elasticity increases by 19% in this period while the evolution of the distance elasticity is best described as flat. It follows that the elasticity of trade costs to distance has decreased by at least 17% in 1970-2009.

What is the economic interpretation of an increasing substitutability parameter measured on aggregate data?

First, the degree of perceived similarity of country-composite goods may have increased. Since the 1960s, a growing number of countries started producing a set of goods similar to that of developed countries. This process has increased the number of available varieties and, potentially, their degree of similarity.[32]

Second, composition effects may have lead to changes in the parameter estimated on aggregate data. If the reduction in trade barriers has led to the expansion in the range of traded

[31] 11 The elasticity decreases by 5% when the evolution of ρ is computed from the ratio of trends, and by 7% when it is computed as the trend of the ratios.

[32] 12 (Schott 2004) documents increased similarity in the set of exported goods of US trade partners while (Broda and Weinstein (2006)) document the increase in the number of imported varieties since the 1970s.

goods, trade in previously non-traded sectors could modify measured substitutability of country-composite goods. Non-uniform reductions in sectoral trade costs would also modify the composition of world trade, leading to a change in the substitutability parameter measured on aggregate data. However, at first approximation, the rising importance of manufactures compared to primary products in world trade should have reduced substituability.

Ideally, we would like to separate out the impact of composition and sector-specific effects to quantify the net effect of increased perceived similarity of country-composite goods. This is however impossible because, as shown by (Imbs et Méjean 2013), the parameter estimated on aggregate data cannot be mimicked by a weighted average of sectoral parameters. The bottom line is that an increase in measured substitutability for country-composite goods is consistent with complex competition dynamics in price and quality documented by (Amiti et Khandelwal 2013) as well as with increased vertical specialization of countries within sectors documented by (Fontagné, Gaulier, et Zignago 2008).

4 Conclusion

The estimated effect of distance in gravity equations has increased in the past fifty years despite substantial innovation in transportation and communication: this is the 'distance puzzle'. Using COMTRADE 4-digit bilateral trade data in 1962-2009, this paper finds that the evolution of the elasticity of trade flows to trade costs, referred to as the 'trade elasticity', provides a direct explanation of the increasing distance elasticity of trade. Increased sensitivity of trade flows to relative prices has more than compensated the reduction in the elasticity of trade costs to distance.

The paper proceeds in three steps. First, it shows that the distance puzzle is a feature of our data by estimating yearly cross-section gravity equations. In the baseline estimation the distance coefficient has increased by 7% from 1962 to 2009. This result holds when we correct for changes in the sample of trading partners and the composition of world trade. Taking into account FTAs seems to solve the distance puzzle, but this might be an artefact of their growing importance: introducing FTAs dummies amounts to adding a time-growing number of proximity controls in the estimation.

Second, the paper suggests an *ad hoc* method of measuring structural heterogeneity in the Armington framework. In the main theoretical foundations of the gravity equation the distance coefficient is the product of the elasticity of trade costs to distance and a measure of heterogeneity.

In the Armington framework, heterogeneity is inter-country and intra-sector. The trade elasticity corresponds to the degree of perceived substitutability of country-specific varieties of each good, which can be approximated by studying the relations between the price level and the market share of importers in destination markets.

Third, the paper estimates the evolution of the trade elasticity in the Armington framework, i.e. the substitution elasticity between country composite goods. It uses 4-digit unit values as proxies for sectoral prices. In our method, unobserved unit values for zero trade flows lead to an overestimation bias that is reduced over time. As the estimated elasticity still increases in absolute value thisevolution provides a lower bound on the increase in the absolute value of the underlying trade elasticity. Once instrumented by bilateral real exchange rates, the estimated elasticity increases by 13% between 1963 and 2009. The evolution of the distance coefficient is thus compatible with a decrease of the elasticity of trade costs to distance of at least 5 to 7%. This reduction in the elasticity of trade costs to distance is even more pronounced if we focus on the period in which air transportation starts playing an important role in bilateral trade. We find that from 1970 to 2009 the elasticity of trade costs to distance has decreased by 17% while the perceived substitutability of countries' product bundles has increased by at least 19%.

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A Full and superbalanced samples

The full sample contains 205 reporters (R) and 227 partners (P) listed in tables 3 and 4 below. 'S' indicates that the country is present in the superbalanced sample.

In the full sample, several countries shift from reporting trade on an individual basis to reporting trade jointly with another country. This is the case of Belgium and Luxembourg, as well as Eritrea and Ethiopia. For consistency, we use a single country identifier for each of these two pairs. A single country identifier is also used for Yougoslavia and for Serbia and Montenegro. Fig.7 shows the distribution of pairs in the full sample according to the number of years in which the pair reports a positive amount of trade.

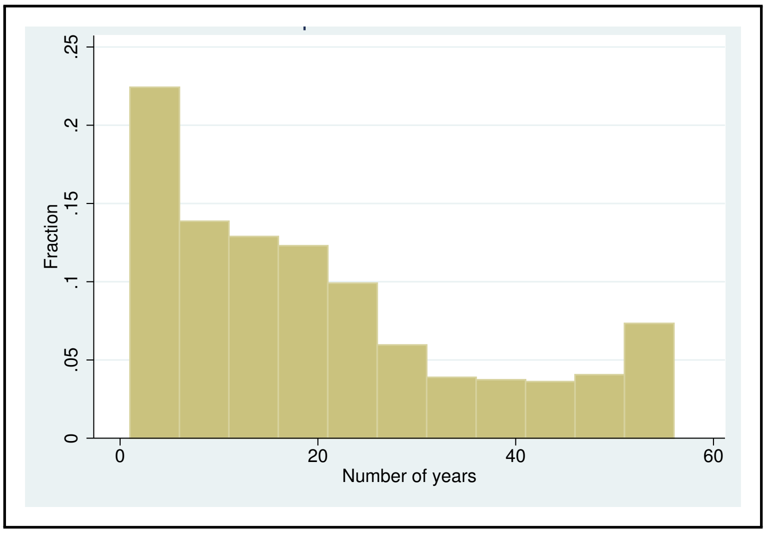

Figure 7: Number of years each pair is present in the sample

The superbalanced sample corresponds to the subsample of pairs which trade both ways in each and every year in 1962-2013. To avoid discarding pairs which fall out of the superbalanced sample because countries split up or reunite at some point in 1962-2013, we introduce several additional single country identifiers before constructing the superbalanced sample. Consequently, Germany is present in the superbalanced sample.[33]

The superbalanced sample comprises 786 trading pairs and corresponds to 32 countries. This is less than the 992 pairs which would be observed if each reporter traded both ways with every other country. Indeed, the set of countries which trade with every other country in each

[33] A single identifier is used for East, West, and reunited Germany. A single identifier is used for the Czech Republic, Slovakia, and Czechoslovakia. A single identifier is used for the USSR and the 15 countries formed after the USSR split up. The 15 countries which constituted the USSR are absent from the superbalanced sample because the USSR is never a reporter to COMTRADE. The Czech Republic and Slovakia also drop out because there is no other country in the sample with which they have two-way trade in every year.

year and which we refer to as the 'square sample' comprises just 21 countries (420 pairs). Trade coverage in the superbalanced sample decreases from 68 to 39% of total trade over 1962-2013 while it is reduced from 54 to 28% for the square sample (see Fig.8).

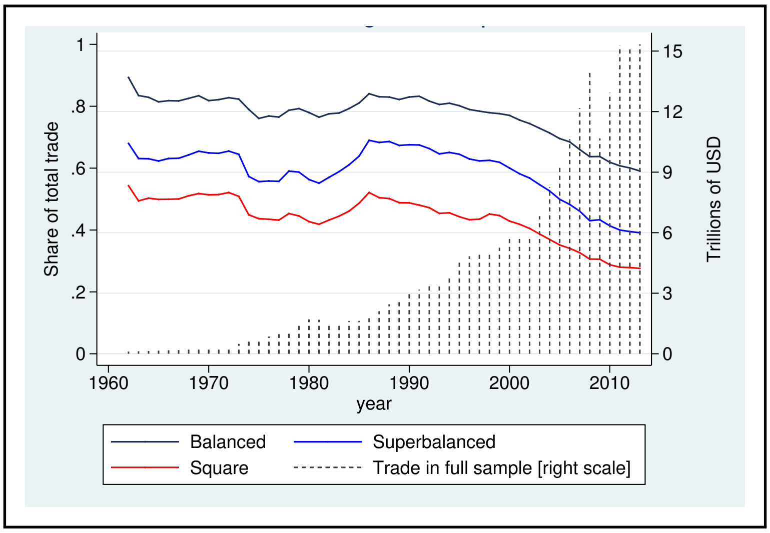

Figure 8: Trade coverage in 1962-2013

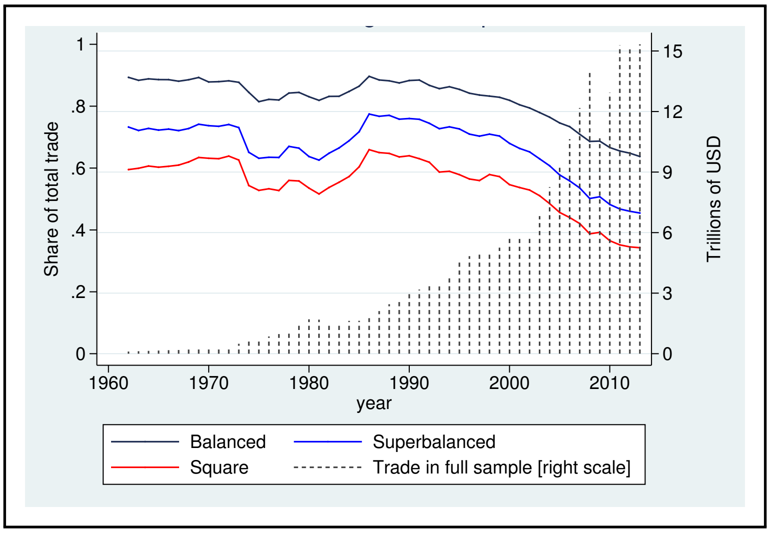

Figure 9: Trade coverage in 1965-2013

To check the sensitivity of trade coverage to the choice of the benchmark year we redefine the set of stable pairs as partners trading both ways in each year in 1965-2013.[34] The superbalanced sample for 1965-2013 contains 1286 pairs that comprise 42 countries (out of 1722 possible pairs). The square sample for 1965-2013 contains 23 countries (506 pairs).

[34] The number of reporters increases from 70 in 1962 to 86 in 1965 to 111 in 1970 and 129 in 1975. But the reduction in trade coverage between the 1970s and the 1990s is qualitatively similar if we choose 1970 or 1975.

As illustrated in Fig.9, the evolution of trade coverage is not sensitive to the choice of the starting year. Albeit from a higher level (73% in 1965), trade coverage drops to 45% by 2013. Indeed, the main reason for the reduction in trade coverage since the mid-1990s is due to the absence of China and of Central and Eastern European countries from the superbalanced sample. These countries drop out because they do not report trade to COMTRADE until the more recent period.[35]

Another way to check the sensitivity of trade coverage to the choice of the starting year is to compute the evolution of trade coverage for the sample of pairs which trade both ways in a specific year. The difference with the superbalanced sample is that we now relax the constraint that the pair have two-way trade in every year in 1962-2013. The share of total annual trade attributable to two-way pairs in 1962 and in 1965 is shown in Fig.10. As previously, the level but not the evolution of trade coverage is affected by the choice of the starting year.

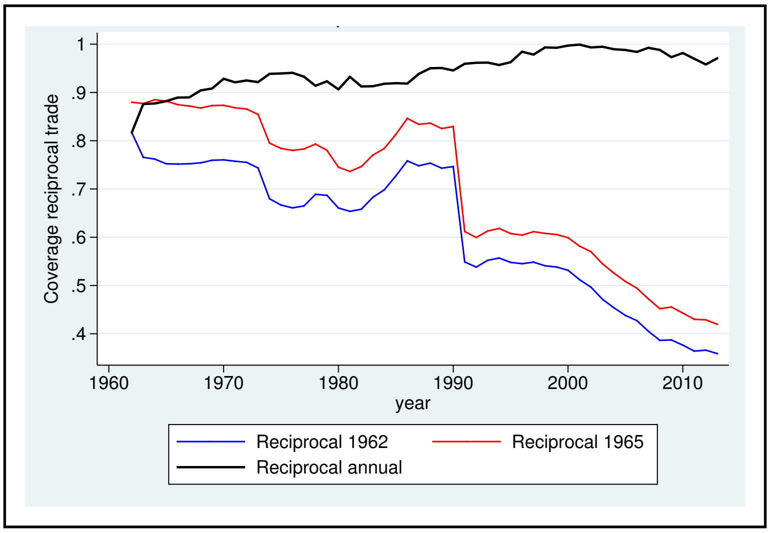

Figure 10: Trade coverage in reciprocal trade

Helpman et al. (2008) document that the enlargement of the set of trading partners did not contribute in a major way to the growth of world trade in 1970-1997 because most of the increase was driven by pairs which traded both ways in 1970. We nuance this finding by showing that since the mid-1990s new reciprocal trade relationships did contribute strongly to the growth of world trade. Fig.10 illustrates that while more than 70% (resp. 80%) of world trade in 1990 is attributable to pairs that traded both ways in 1962 (resp. 1965), less than 40% (resp. 50%) is still attributable to such pairs in 2013.

[35] For example, China reports trade to UN COMTRADE in 1987-2013.

We document that these new trade relationships were formed between countries trading both ways. To illustrate, Fig.10 also shows the share of total trade attributable to pairs which trade both ways in a given year (in black). Since the 1990s more than 95% of total annual trade takes place between pairs which trade both ways. One-way trade flows represent a marginal and decreasing share of world trade. The greater frequency of one-way trade relationships observed in the early years of the sample may be in part attributable to the fact that the number of countries reporting trade to COMTRADE was initially relatively small. Consequently, reported zeros may be at least in part statistical zeros, i.e. non-zero trade flows reported as zeros due to missing reports to COMTRADE.

Table 3: List of countries in the full and superbalanced samples

Country name	Status	Country name	Status	Country name	Status
Afghanistan	*R;P*	French Polynesia	*R;P*	N. Mariana Islands	P
Albania	*R;P*	French S. Antartic terr.	P	Norway	*R;P*
Algeria	*R;P*	Gabon	*R;P*	Oman	*R;P*
Andorra	*R;P*	Gambia	*R;P*	Pakistan	*R;P*
Angola	*R;P*	Georgia	*R;P*	Palau	P
Anguilla	*R;P*	**Germany**	**R;P;S**	Panama (Fm Panama Cz)	*R;P*
Antigua-Barbuda	*R;P*	Ghana	*R;P*	Papua New Guinea	*R;P*
Argentina	**R;P;S**	Gibraltar	P	**Paraguay**	**R;P;S**
Armenia	*R;P*	**Greece**	**R;P;S**	Peru	*R;P*
Aruba	*R;P*	Greenland	*R;P*	**Philippines**	**R;P;S**
Australia	R;P	Grenada	*R;P*	Pitcairn	P
Austria	R;P	Guadeloupe	*R;P*	Poland	*R;P*
Azerbaijan	*R;P*	Guatemala	*R;P*	**Portugal**	**R;P;S**
Bahamas	*R;P*	Guinea	*R;P*	Qatar	*R;P*
Bahrain	*R;P*	Guinea-Bissau	*R;P*	Reunion	*R;P*
Bangladesh	*R;P*	Guyana	*R;P*	Romania	*R;P*
Barbados	*R;P*	Haiti	*R;P*	Russian Federation	*R;P*
Belarus	*R;P*	Honduras	*R;P*	Rwanda	*R;P*
Belgium-Luxembourg	**R;P;S**	**Hong Kong**	**R;P;S**	St. Helena	P
Belize	*R;P*	Hungary	*R;P*	St. Kitts and Nevis	*R;P*
Benin	*R;P*	**Iceland**	**R;P;S**	St. Lucia	*R;P*
Bermuda	*R;P*	India	*R;P*	St. Vincent-Grenadines	*R;P*
Bhutan	*R;P*	Indonesia	*R;P*	Samoa	*R;P*
Bolivia	*R;P*	Iran	*R;P*	San Marino	P
Bosnia-Herzeg.	*R;P*	Iraq	*R;P*	Sao Tome-Principe	*R;P*
Botswana	*R;P*	Ireland	R;P	Saudi Arabia	*R;P*
Brazil	**R;P;S**	**Israel**	**R;P;S**	Senegal	*R;P*
Br. Virgin Islands	P	**Italy**	**R;P;S**	Serbia-Montenegro	*R;P*
Brunei Darussalam	*R;P*	Jamaica	*R;P*	Seychelles	*R;P*
Bulgaria	*R;P*	**Japan**	**R;P;S**	Sierra Leone	*R;P*
Burkina Faso	*R;P*	Jordan	*R;P*	**Singapore**	**R;P;S**
Burma (Myanmar)	*R;P*	Kazakstan	*R;P*	Slovakia	*R;P*
Burundi	*R;P*	Kenya	*R;P*	Slovenia	*R;P*
Cambodia	*R;P*	Kiribati	*R;P*	Solomon Islands	*R;P*
Cameroon	*R;P*	**Korea**	**R;P;S**	Somalia	*R;P*
Canada	**R;P;S**	DPR of Korea	P	South Africa	*R;P*

Table 4: List of countries in the full and superbalanced samples: Contd.

Country name	**Status**	**Country name**	**Status**	**Country name**	**Status**
Cape Verde	*R;P*	Kuwait	*R;P*	Soviet Union	P
Cayman Islands	P	Kyrgyzstan	*R;P*	**Spain**	**R;P;S**
C.African Republic	*R;P*	Lao PDR	*R;P*	Sri Lanka	*R;P*
Chad	*R;P*	Latvia	*R;P*	St. Pierre and Miquelon	*R;P*
Chile	**R;P;S**	Lebanon	*R;P*	Sudan	*R;P*
China	*R;P*	Lesotho	*R;P*	Suriname	*R;P*
Christmas Island	P	Liberia	*R;P*	Swaziland	*R;P*
Cocos Islands	P	Libya	*R;P*	**Sweden**	**R;P;S**
Colombia	**R;P;S**	Lithuania	*R;P*	**Switzerland**	**R;P;S**
Comoros	*R;P*	Luxembourg	R;P	Syria	*R;P*
Congo	*R;P*	Macau (Aomen)	*R;P*		
Dem. Rep. of Congo	*R;P*	Macedonia	*R;P*	Tajikistan	*R;P*
Cook Islands	*R;P*	Madagascar	*R;P*	Tanzania	*R;P*
Costa Rica	*R;P*	Malawi	*R;P*	**Thailand**	**R;P;S**
Croatia	*R;P*	**Malaysia**	**R;P;S**	Togo	*R;P*
Cuba	*R;P*	Maldives	*R;P*	Tokelau	P
Cyprus	*R;P*	Mali	*R;P*	Tonga	*R;P*
Czech Republic	*R;P*	Malta	*R;P*	Trinidad-Tobago	*R;P*
Czechoslovakia	*R;P*	Marshall Islands	P	**Tunisia**	**R;P;S**
Côte d'Ivoire	*R;P*	Martinique	*R;P*	**Turkey**	**R;P;S**
Denmark	**R;P;S**	Mauritania	*R;P*	Turkmenistan	*R;P*
Djibouti	*R;P*	Mauritius	*R;P*	Turks-Caicos Islands	*R;P*
Dominica	*R;P*	**Mexico**	**R;P;S**	Tuvalu	*R;P*
Dominican Republic	*R;P*	Micronesia	P	Uganda	*R;P*
East Germany (DDR)	R;P	Moldova	*R;P*	Ukraine	*R;P*
East Timor	*R;P*	Mongolia	*R;P*	United Arab Emirates	*R;P*
Ecuador	*R;P*	Montserrat	*R;P*	**United Kingdom**	**R;P;S**
Egypt	*R;P*	Morocco	*R;P*	**USA**	**R;P;S**
El Salvador	R;P	Mozambique	*R;P*	Uruguay	*R;P*
Equatorial Guinea	P	Namibia	*R;P*	Uzbekistan	P
Eritrea	*R;P*	Nauru	P	Vanuatu	*R;P*
Estonia	*R;P*	Nepal	*R;P*	**Venezuela**	**R;P;S**
Ethiopia	*R;P*	Netherland Antilles	*R;P*	Vietnam (Fm Vietnam Rp)	*R;P*
Falkland Islands	P	**Netherlands**	**R;P;S**	Wallis-Futuna	*R;P*
Palestine	*R;P*	New Caledonia	*R;P*	West Germany (FRG)	R;P
Fiji	*R;P*	New Zealand	*R;P*	Western Sahara	P
Finland	R;P	Nicaragua	*R;P*	Yemen	*R;P*
France	**R;P;S**	Niger	*R;P*	Yugoslavia (Serbia-Mont.)	*R;P*
French Guiana	*R;P*	Nigeria	*R;P*	Zambia	*R;P*
Niue	P	Norfolk Island	P	Zimbabwe	*R;P*