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Return of the Tariffs: The Interwar Trade Collapse Revisited

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Abstract

Was the collapse of world trade between 1928 and 1937 caused by higher transport costs, increased protectionism or the collapse of the gold standard? Using recent advances in the estimation of gravity equations, I examine the partial and general equilibrium effects of bilateral distance, international borders, and the payment system on trade. My results suggest that had average tariff and non-tariff trade barriers remained at their 1928 level, total international trade would have been 64.6% higher in 1937. Had the gold standard not collapsed in 1931 and had the British Empire not departed to establish its own currency and trade blocs, international trade would have been 3% larger. Finally, had transport costs remained at their 1928 level, global trade would not have been significantly different nine years on. These results are supported by over 6,000 new hand-collected observations of ad-valorem ocean freight rates for cotton, which show an average increase of only 1.2 percentage points between 1928 and 1936. When expressed as an index, the movement of freight rates mirrors the evolution of the elasticity of trade to distance over the period.

Keywords: International trade; Gravity equation; Great Depression

JEL Classification: F10; F12; F13; F33; N70

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Contents

1 The Tariff Menace

In 2016, the Brexit referendum and the U.S. presidential election of Donald Trump indicated that the globalization of the last decades might be coming to an end. These fears seemed to be confirmed in 2018, when the U.S. imposed tariffs on solar panels, aluminum, steel, and washing machines from most countries including its closest trading partners. These tariffs caused retaliation from many countries and fears are mounting that U.S. - China relations might escalate into a large-scale trade war. The increased trade tensions and the threat of a shift away from a multilateral trading system forced the International Monetary Fund (IMF) to revise downwards its World Economic Outlook for 2019. In the eyes of the IMF (2018) the "intensification of trade tensions, and the associated rise in policy uncertainty, could dent business and financial market sentiment, trigger financial market volatility, and slow investment and trade. Higher trade barriers would disrupt global supply chains, [...] lowering global productivity and [harming] low-income households disproportionately."

Many observers point to the risk of returning to a 1930s beggar-thy-neighbor trade policy.[1] Strikingly, anti-globalization rhetoric and policies have increased only recently even though global trade has already been slowing down since 2012. Because protectionist policies were not put in place after the disruption of the 2008 financial crisis, economists were initially looking for other explanations for the stall in world trade growth. Economists looked at the dramatic collapse of world trade in the 1930s to draw conclusions about the causes of deglobalization. Estevadeordal et al. (2003) explore the causes of the collapse of world trade during the interwar period. They argue that because productivity growth in the shipping sector was slower than average total factor productivity (TFP) growth, real transport costs rose in the interwar period. To explain the contemporary trade stagnation, Krugman (2016) popularized Estevadeordal et al.'s (2003) interpretation, arguing that higher transport costs were the main determinant in the collapse of world trade in the 1930s.

Independently of whether one can apply this argument to the present period of trade stagnation, it is worth looking at the interwar period again, which has historically been associated with rising tariff and non-tariff trade barriers rather than with rising transport costs. Estevadeordal et al. (2003) present increased protectionism, the collapse of the international payment system and rising transport costs as possible causes of the collapse of world trade in the 1930s. This paper takes a fresh look at the three candidates. It makes use of recent advances in the workhorse model of international trade, the gravity model that seeks to explain bilateral trade flows. Changes in the partial effects of distance, borders and the payment system are estimated for a sample of 36 countries and the resulting trade cost function is incorporated into a full endowment general equilibrium (GE) model, which allows me to revisit the horse race of Estevadeordal et al. (2003).

My regression results suggest that the border effect, which measures the thickness of international borders and serves as a proxy for average bilateral tariff and non-tariff trade barriers, had increased by 89% between 1925 and 1937. At the same time, there was no significant change in the trade-reducing effect of bilateral distance. The trade-increasing effect of the gold standard was relatively small. Being on gold increased members' bilateral trade by only 9% and this effect is relativized when taking into account the potentially trade-increasing effects of the trade and currency blocs that

[1] See for example Stephens (2018)

followed the collapse of the gold standard. Indeed, my gravity model, which controls for the average effect of de-globalization, provides evidence that one trade bloc, the imperial preference system (IPS), and one currency bloc, the sterling bloc, increased bilateral trade among their members. This contrasts with previous studies, which find that no trade or currency bloc in the 1930s increased trade. The finding that the trade-increasing effects of the IPS and the sterling bloc are large, positive, significant and free from reverse causality has an important implication: The economic benefits that could be reaped from a retreat into empire, made Britain's exit from the multilateral trade and payment system, the gold standard, less severe than for other countries.

My specification of the structural gravity model also serves to explore the seemingly paradoxical result of a declining elasticity of trade to distance that previous studies have found, despite rising transport costs during the interwar period. My results support previous explanations for the distance puzzle, which argue that commercial and financial policies increasingly dominated the effect of distance, but also stress the heterogeneity of tariff rates between trading partners. The "simple" solution to the post-war distance puzzle proposed in the international trade literature is not sufficient to solve the interwar distance puzzle. Instead, it requires the inclusion of a large and significant effect of the IPS, indicating preferential tariff rates between the British Empire and its dominions, to get an unbiased coefficient of distance as a trade cost factor. Only after the inclusion of the empire effect can transport costs be reasonably proxied by the distance elasticity, suggesting that the IPS, agreed upon at the Ottawa conference in 1932, was successful in defying gravity, i.e. making the physical trade cost of distance relatively less important.

The estimated partial effects yield the trade cost function, which is incorporated into a full endowment GE model. Analyzing three different counterfactual scenarios in the GE model is akin to Estevadeordal et al.'s (2003) horse race and allows me to answer the following questions:

- What would have been the level of world trade in 1937, had the gold exchange standard not collapsed into a system of trade and currency blocs?
- What would have been the level of world trade in 1937, had tariff and non-tariff trade barriers remained at their 1928 levels?
- What would have been the level of world trade in 1937, had transport costs remained at their 1928 level?

The results suggest that if transport costs had remained at their 1928 level, trade in 1937 would have been 19% lower. This effect is however not significantly different from zero. Had the gold standard not collapsed and had the IPS and the sterling bloc not formed, global trade would have been merely 3% larger in 1937. But had average tariff and non-tariff trade barriers, as proxied by the border effect, not increased after 1928, world trade would have been 64% larger in 1937. These results contrast sharply with the results of Estevadeordal et al. (2003), the only study to date that quantifies the individual contributions of tariffs, transport costs and the payment system. My results provide quantitative estimates that reestablish the conventional narrative that protectionism was the culprit of the interwar trade bust.

To support my results, I present additional evidence on real transport costs during the interwar period. I manually collected high-frequency data on cotton freight rates from New York along 21 routes from 1925 to 1936, which I deflate by the product price in New York to get ad-valorem freight rates. The data imply only a marginal increase in ad-valorem freight rates. On average, real freight rates increased by less than two percentage points between 1925 and 1936. This lends support to the regression results I obtain from estimating the gravity model.

Finally, I match the real freight rates with the quantities shipped to the destinations. This allows me to create a Laspeyres index, which is an additional contribution to the interwar shipping literature, which until today had relied exclusively on the Isserlis-index. Contrary to the Isserlis-Index, my new cotton freight index covers different shipping routes and is based on freight rates for the liner industry. It tracks historical events such as the coal strike of 1926 well and shows an increase of 50% between 1925 and 1936. More importantly, my transport cost index mirrors the movement of the distance elasticity from the gravity model. Both, the index and the distance elasticity, shoot up during the Great Deflation of 1929-1933, which I attribute to cartelization, rather than to a productivity slowdown in the shipping sector. Future research could delve into the question of the role that the sharp increase in real transport costs around 1931 played in the initial trade bust.

2 Gravity Between the Wars: The Empire Adrift

For international trade economists, the contraction of world trade during the first phase of the Great Depression is remarkable, both in absolute and relative terms to GDP. From 1929 to 1933, world exports in constant prices fell by 35%. When output started to recover, it was not followed by international trade, and in 1937 real volume of world trade was barely 95% of its 1928 level.[2] What caused this collapse in international trade?

A major factor of the trade bust was the fall in world income, but again this cannot explain the low level of trade after income had recovered. The period was also marked by a surge in protectionism following the infamous Smoot-Hawley Tariff imposed by the U.S. in 1930. But tariffs are only one of several factors that increased the costs of trade between countries. The financial crisis in continental Europe in summer of 1931 marked the beginning of the collapse of the gold standard and was followed by devaluation, the introduction of capital controls and the formation of new currency blocs. This collapse of the multilateral payment system is seen as an important factor in its own right, separately from tariff and non-tariff trade barriers. A third potential factor is an increase in real transportation costs (Irwin, 2011).

Estevadeordal et al. (2003) innovatively investigated the relative impact of these three factors by using a gravity model.[3] Of the six percentage point decline in the trade-to-GDP ratio between 1929 and 1938 they attribute 29% to the collapse of the payment system, 27% to higher transport costs and only 14% to higher tariffs. Arguing that the interwar trade collapse was caused more by higher shipping costs than by rising tariffs earned them a reputation as revisionists (Jacks et al., 2011).

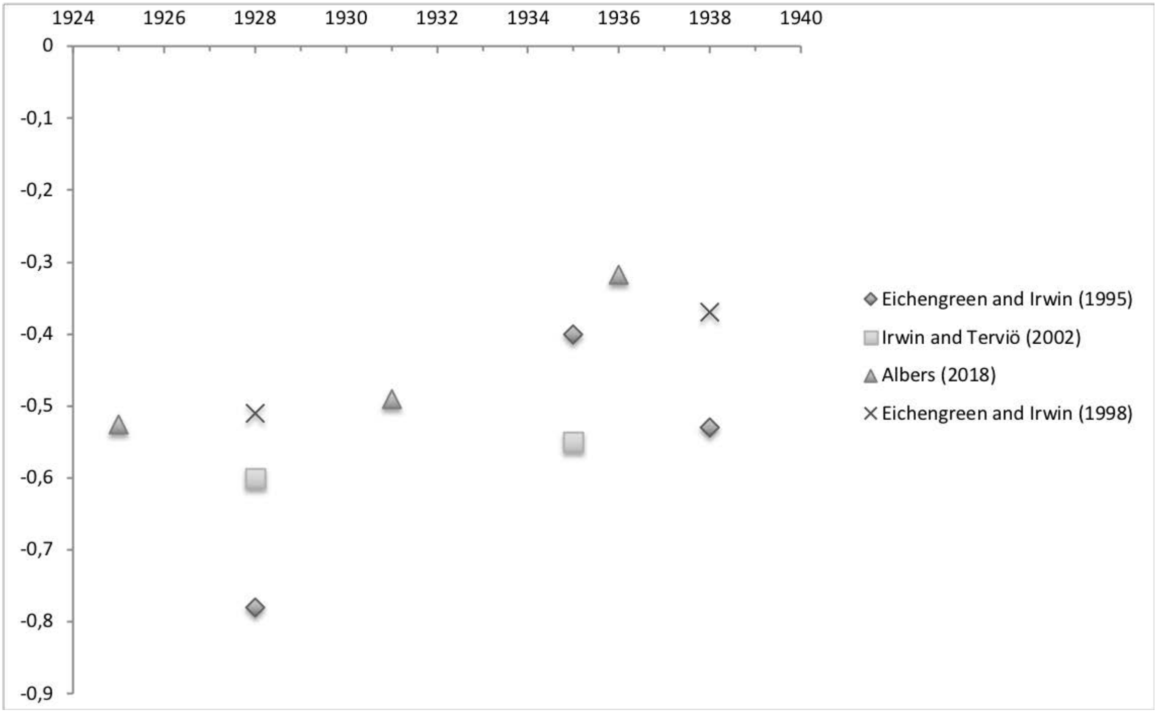
Albers (2018) challenges Estevadeordal et al.'s (2003) finding by estimating a gravity model for twelve consecutive years from 1925 to 1936 to determine the elasticity of trade with respect to distance. He confirms an earlier finding by Eichengreen and Irwin (1995) of a rise in the coefficient of distance from 1929 onwards. Since the distance elasticity is negative, a rise (decrease in absolute value) means that distance becomes less important as the world enters the depression phase. Because distance becomes less important at a time when real transport costs are rising, Albers (2018) names his finding the "interwar distance puzzle" referring to the postwar distance puzzle in the meta-study of Disdier and Head (2008). Albers (2018) considers that his finding means that tariffs were becoming a more important factor in determining trade relative to transport costs. Since the effect of tariffs outweighs the effect of transport costs, the relative importance of distance diminishes. This is in line with the interpretation of Eichengreen and Irwin (1995) who argue, "that commercial and financial policies increasingly dominated the effects of geography". Following Disdier and Head (2008), I analyze all studies that provide distance coefficients for individual years of the interwar period. Figure (1) plots the coefficients of the main regression of each study over time. We indeed observe a significant decline in the absolute value of the distance-coefficient.

Another recent study by Fouquin and Hugot (2016a) estimates the distance elasticity as yearly

[2] Statistics in this paragraph are computed from table D.14 in Federico and Tena Junguito (2016), which shows world exports in constant 1913 USD.

[3] Another study that quantitatively investigates the trade collapse is Madsen (2001), who argues that approximately 41% of the world trade collapse over the period 1929 to 1932 can be attributed to tariff and non-tariff trade barriers, and the rest is due to declining incomes. However, Madsen (2001) only deals with the immediate depression period and does not consider transport costs or the collapse of the payment system as possible causes.

Figure 1: The Interwar Distance Elasticity Puzzle



Notes: This figure plots the distance coefficients of four studies (Albers, 2018; Eichengreen and Irwin, 1995, 1998; Irwin and Terviö, 2002) over time.

repeated cross sections from 1827 to 2014. While their paper does not present any numerical results, their figure 18 shows that the distance elasticity has roughly the same value in 1938 as in 1920. This stands in stark contrast to the studies in figure (1).[4]

Recent research has solved the postwar distance puzzle by incorporating internal trade and internal distance into the gravity model. Yotov (2012) argues that the puzzle had persisted because previous studies estimated *international* trade costs relative to other *international* trade costs, when one should instead measure *international* relative to *intranational* trade costs. A bias in the distance coefficient could arise, for example, if a country unilaterally increased its tariffs. Nearby economies previously exporting to this country would redirect some of their exports to more distant countries. A regression that does not include intranational trade then sees the effect of international distance decline.[5]

[4]Unfortunately, the estimation strategy by Fouquin and Hugot (2016a) is not completely clear. For example, they do not seem to use a balanced sample, but instead have a different number of observations for each year. However, Fouquin and Hugot (2016a) include internal trade and the border effect in their regression, which is part of the solution to the distance puzzle as discussed below.

[5]Where trade flows are redirected to and how strong this bias is depends on exporters' and importers' trade openness, their economic size and market integration, or their location in the world economy (see e.g. the discussion in Liu and Meissner (2015)). Moreover, a change in trade policy or transportation costs affects countries differently depending on their product mix (see appendix (A.1)). Finally, variations in productivity lead to variations in the extensive margin. If a change in trade policy or transport costs affects the fixed costs of exporting, heterogenous effects arise in the extensive margin of trade (Chaney, 2008). Such heterogeneity across countries of a changing distance elasticity has also been observed by Borchert and Yotov (2017) who suggest the inclusion of a set of country-specific fixed effects for internal trade. As discussed below, I deal with such problems by adopting the estimation strategy of Bergstrand et al. (2015).

Internal trade is also included in Bergstrand et al. (2015) who estimate a panel version of the gravity model from 1990 to 2002. Their model allows to control for unobserved heterogeneity across country pairs by including a set of country-pair fixed effects. They also include a dummy variable for international borders that captures the average decrease (or increase) in tariff and non-tariff trade barriers or more generally the average effect of globalization (or deglobalization). Because Bergstrand et al. (2015) provide the best solution for the post-war distance puzzle to date, I follow their econometric strategy to measure changes in the distance elasticity in the interwar period. This approach should provide us with a good proxy for real transportation costs. Still, I find that the distance elasticity has fallen over the course of the 1930s even after controlling for international borders.

The state-of-the-art features of the gravity model are also indispensable to get an unbiased estimate for dummy variables that measure the effect of trade agreements, currency unions or other *economic integration areas* (Bergstrand et al., 2015). Including these specifications is an improvement over earlier studies of the interwar period such as Gowa and Hicks (2013), who investigate the system of trade and currency blocs that evolved after the collapse of the gold standard, or Jacks (2014), who asks whether Canada was able to divert trade flows towards members of the IPS agreed upon at the Ottawa conference in 1932. Gowa and Hicks (2013) find that no bloc increased or decreased trade among its members, and also Jacks (2014) finds that Canada was not able to defy gravity and divert trade flows towards members of the IPS. Because these studies do not include internal trade, they are not able to account for the effect of deglobalization which increases internal trade relative to international trade. The absence of accounting for this effect biases the estimate for trade and currency blocs downwards. Controlling for deglobalization, I find that the IPS is economically and statistically significant. Moreover, there is enough evidence to suggest that sterling bloc membership increased trade between members. Including these two blocs is essential to get an unbiased estimate of distance that proxies transport costs. These results lend support to a recent study by De Bromhead et al. (2019) who argue that the discriminatory trade policies of the British Empire and its dominions were a significant factor in shifting trade towards the empire. In that sense, it is no surprise that the interwar distance puzzle is resolved once we control for the IPS and sterling bloc. Following the Ottawa conference in 1932, the British Empire decoupled itself and drifted away from the rest of the world. The breakaway of the empire from the rest of the world effectively decreased the relative distance between the empire and its dominions, while decreasing the relative distance between countries of the rest of the world. Overall, the preferential tariff rates and the many exceptions granted to the dominions on non-tariff barriers rendered distance relatively less important as a trade cost.

Equipped with a complete trade cost function, I conduct GE static exercises to analyze the individual contributions of transport costs, the multilateral trade and payment systems, and tariff and non-tariff trade barriers to the global trade bust.

This paper contributes to three strands of the literature. First, my results are a vindication of Bergstrand et al.'s (2015) proposed solution to the distance elasticity puzzle, which goes beyond Yotov's (2012) inclusion of intranational trade and suggests that ignoring the effects of *economic integration areas*, or blocs, biases the coefficient on distance. Second, it contributes to the debate on the system of trade and currency blocs in the 1930s by showing that the IPS and the sterling bloc were successful in shifting trade towards their members. Third, and most importantly, my results question the existence of a rise in real transport costs and challenge the findings of Estevadeordal et al. (2003). This study supports the more conventional explanation for the interwar trade collapse: protectionism.

3 Interwar Gravity Redux

This section first deals with the theoretical foundation of the structural gravity model and then uses the model to estimate the direct, or partial, effects of distance, international borders and the payment system on international trade. Finally, I use the resulting trade cost function to perform GE analysis in three counterfactual scenarios.

3.1 Methodology and Data

Since the 1960s, trade economists have used the "gravity equation" to provide econometric estimates for the effects of distance, national borders, currency unions, and other measures of trade costs on bilateral international trade flows. It is only in the last two decades that the gravity model has evolved from a simplistic analogy with Newtonian physics to the workhorse model of international trade.

The early 2000s saw a gravity revolution caused by the influential works of Eaton and Kortum (2002) and Anderson and Van Wincoop [AvW] (2003), who endowed the gravity equation with micro-foundations. Eaton and Kortum (2002) derive the gravity equation from a Ricardian supply-side framework, while AvW derive the gravity model from a demand-side Armington (i.e. CES-National Product Differentiation) framework. Although the starting points of these two studies are radically different, they arrive almost at the same results. Indeed, Arkolakis et al. (2012) have shown that the gravity equation can be derived not only from an Armington and Ricardian framework but from an even wider range of trade models, including models in the spirit of Krugman (1980) and Melitz (2003). The present study uses the Armington framework in the tradition of AvW to analyze the impact of different trade costs on world trade during the interwar period. However, Allen et al. (2014) have recently developed a universal gravity framework with sufficient conditions for the existence and uniqueness of the trade equilibrium for a wide class of GE models including AvW. Therefore, the macroeconomic conditions inherent in the gravity trade model impose sufficient structure so that its particular microeconomic details do not pose a problem in its characterization.

The model considered in this study consists of N countries, where each country produces a variety of goods that is traded with all other countries. Denoting the fixed supply of each good with Q_i and the factory-gate price with p_i, the value of production, or income, in country i is defined as $Y_i = p_i Q_i$. Aggregate expenditure is defined as $E_i = \phi_i Y_i$, where ϕ_i is an exogenous parameter defining the relation between the value of output and aggregate expenditure, such that when $\phi_i > 1$, country i faces a trade deficit, while country i runs a trade surplus when $1 > \phi_i > 0$. The complete gravity model that explains exports (X_{ij}) from country i to j is described in equations (3.1) to (3.5).[6] Π_i and P_j are structural terms which AvW call outward and inward multilateral resistance terms. τ_{ij} is the trade cost factor between i and j, σ is the elasticity of substitution and α_i is the CES preference parameter.

[6] For a full derivation of the micro-founded gravity model from an Armington framework the reader is directed to AvW's original article or one of the guides and handbooks on the topic. The most recent guide on the gravity model on which the present study draws extensively is Yotov et al. (2016). Equations (3.1 - 3.5) and the estimation procedure (including much of the Stata code) for the GE analysis are adapted from Yotov et al. (2016) and can be downloaded at https://vi.unctad.org/tpa/web/vol2/vol2home.html

$$\begin{array}{c}\text{Full}\\\text{Endowment}\\\text{(GE)}\end{array}\begin{cases}\begin{array}{c}\text{Conditional}\\\text{(GE)}\end{array}\begin{cases}\text{Direct (PE)}\begin{cases}X_{ij} = Y_i E_j \left(\dfrac{\tau_{ij}}{P_j \Pi_i}\right)^{1-\sigma} & (3.1)\end{cases}\\[2ex] \Pi_i \equiv \left(\displaystyle\sum_{j=1}^{C} E_j \left(\dfrac{\tau_{ij}}{P_j}\right)^{1-\sigma}\right)^{1/(1-\sigma)} \quad (3.2)\\[2ex] P_j = \left(\displaystyle\sum_{i=1}^{C} Y_i \left(\dfrac{\tau_{ij}}{\Pi_i}\right)^{1-\sigma}\right)^{1/(1-\sigma)} \quad (3.3)\end{cases}\\[2ex] p_i = (Y_i)^{1/(1-\sigma)} \dfrac{1}{\alpha_i \Pi_i} \quad (3.4)\\[2ex] E_i = \phi_i Y_i = \phi_i p_i Q_i \quad (3.5)\end{cases}$$

Equation (3.1) represents the theoretical gravity equation that governs bilateral trade flows and consists of a size term $Y_i E_j$ and a trade cost term $(\tau_{ij}/P_j \Pi_i)^{1-\sigma}$. At the heart of the structural gravity model are the multilateral resistance terms P_j and Π_i, AvW's key innovation, that differentiates the theory-founded gravity models from the earlier ones. These remoteness terms, which represent the importer j's and exporter i's ease of market access, have to be controlled for to get an unbiased estimate of the partial effect of any factor within the trade cost function when estimating the gravity model econometrically. The theory-founded gravity model then includes trade with all N trading partners of country i including country i itself. Not including intranational trade will result in biased estimates of any partial effect of trade costs since it ignores the effects of trade diversion (Bergstrand et al., 2015). These trade diversion effects work through the multilateral resistance terms and arise because the more integrated country i is with a particular trading partner j, the more remote it becomes relative to all other countries. Because previous studies of the interwar period did not incorporate internal trade, this is the first study to estimate a properly specified theoretical gravity model for the interwar period.

Moreover, any counterfactual analysis, that is performed to examine a change in trade costs between i and j using the partial effect only, as in Estevadeordal et al. (2003), ignores feedback effects affecting other countries. This drawback can be overcome using the GE analysis framework operating via the multilateral resistance channels, captured by equations (3.2) and (3.3). Whereas the partial effect is captured by adjusting bilateral trade costs τ_{ij} while keeping output, expenditure and multilateral resistance terms constant, the *conditional GE* effects allow for adjustment in the multilateral resistance terms.

The channel described in equations (3.4) and (3.5) endogenizes the value of output and expenditure by allowing factory-gate prices to respond to trade cost changes. Analyzing a change in trade costs, the *full endowment GE* then takes into account the associated feedback effects in multilateral resistances, via equation (3.4), and then translates the changes in factory-gate prices into changes in the value of domestic production and aggregate expenditure, via equation (3.5).[7]

[7] An implicit assumption in this paper, as in all standard gravity models, is that the trade cost function is exogenous

In the following subsections, the structural gravity model will be used to evaluate the impact of changes in transport costs, the payment system and political trade barriers on world trade. Section (3.2) makes use of recent econometric advances to estimate the partial effects of distance, international borders and the payment system. Equipped with a complete trade cost function τ_{ij}, section (3.3) then solves the complete GE model in equations (3.1) to (3.5) for the year 1937 and compares it with three counterfactual scenarios (CF 1 - 3) in which specific trade costs are assumed to have remained at their 1928 level. This will answer the following questions:

CF 1: What would have been the level of world trade in 1937, had the gold exchange standard not collapsed into a system of trade and currency blocs?

CF 2: What would have been the level of world trade in 1937, had tariff and non-tariff trade barriers remained at their 1928 levels?

CF 3: What would have been the level of world trade in 1937, had transport costs remained at their 1928 level?

Performing a complete GE analysis representative of world trade requires data on bilateral trade, geographical variables such as distance, gold standard and bloc membership, and internal trade for a large number of countries. To construct the dataset, I draw on a number of existing data sources, the most important of which is Fouquin and Hugot (2016b). Although the past two decades have seen substantial improvements in data for the interwar period, obvious data limitations are still present. The primary difficulty is that for the model to be closed it requires $N * N$ observations per year (i.e. a quadratic matrix of trade relationships between the N trading partners). This presents the risk of missing observations, if one does not want to reduce the number of countries to a non-representative sample with a geographical bias.

The sample used to estimate the trade cost function to perform the GE analysis in section (3.3) consists of 36 countries over five interval years and consequently 6480 observations, of which 881 missing observations are assumed to be zero. The total number of observations in the *GE sample* that take the value zero is 1038. This is a large number of zeros and some country-pairs do not report a single positive trade flow for any year. This causes 390 observations to be dropped from the estimation and in order to get the baseline trade cost function for these country pairs, I apply the two-step procedure suggested by Anderson and Yotov (2016). Fortunately, the restriction of $N * N$ observations can be abandoned in the estimation of the partial effects. This allows me to estimate the trade cost function with more confidence. The estimation of the trade cost function is robust to using the more rigorous *partial sample*.

Three variables are used to estimate the partial effects of transport costs, tariff and non-tariff trade barriers, and the payment system: distance, international borders and gold standard or currency bloc membership. Some authors have argued that distance and borders hinder trade much more than transports costs or tariffs can explain. In particular, Grossman (1998) and Head and Mayer (2013) have argued that distance and borders measure lack of information and home-variety

to income and trade growth. If a negative income shock causes a rise in tariffs, capital controls or other trade barriers, then the role of trade costs in explaining the fall of world trade could, of course, be weaker. Unfortunately, this is an issue that remains outside the scope of this paper.

biased preferences. Here I will make the reasonable assumption that these factors did not change over the period under study so that any change in the elasticity of distance and the border effect can directly be interpreted as a change in the trade decreasing effects of transportation costs and tariff and non-tariff trade barriers.

To determine the partial effect of the collapse of the payment system Estevadeordal et al. (2003) use an indicator variable that describes gold standard adherence between two trading partners. This may be overly pessimistic, as I will show, since the gold standard did not collapse into N national payment systems, but into a system of trade and currency blocs. Contrary to previous studies that estimated the impact of these blocs, I find that membership of the IPS had a large, statistically significant, positive effect on trade. Moreover, I provide evidence that sterling bloc membership increased trade. Controlling for these two blocs in the gravity equation solves the interwar distance puzzle since it removes the omitted variable bias that stems from the heterogeneity in tariff rates. After the Ottawa conference, and as a result of preferential tariff rates, it may have been cheaper for the UK to import goods from far away dominions than from nearby European countries. Since this is based on the presumption that currency blocs did in fact increase trade between members, I will provide additional estimates on the partial effects of all trade and currency blocs and test these blocs for reverse causality.

The last factor in the trade cost function, the border effect, itself will then capture the remaining international trade costs independent of distance, the payment system and other standard gravity control variables such as common language, colonial linkage or contiguity. The border effect should be interpreted as capturing tariff and non-tariff barriers, such as capital controls, quota systems, restrictions on the use of imported inputs by domestic producers, undue controls at frontiers, and regulation.

The complete construction of the data, its sources, and sample selection are discussed in detail in the appendix (A.1).

3.2 Partial Effects

Recent econometric advances in the estimation of the gravity model, discussed in section (2), provide us with reliable estimates of the partial effects of distance, international borders, and the payment system. Using a panel of 36 countries and five years (1925, 1928, 1931, 1934, 1937), I estimate the following equation using the PPML estimator:[8]

$$\begin{aligned} X_{ij,t} = exp[\sum_{T=1928}^{1937} \beta_{1,T} \ln(Dist_{ij,T}) + \sum_{T=1928}^{1937} \beta_{2,T} INTL\_BRDR_{ij,T}]* \\ exp[\beta_3 Cbloc + \beta_4 Cbloc_{t-s} + \beta_5 Cbloc_{t+4} + \gamma_{i,t} + \delta_{j,t} + \phi_{ij}] + \epsilon_{ij,t} \end{aligned} \tag{3.6}$$

The estimation strategy described in (3.6) follows Bergstrand et al. (2015) and estimates a panel

[8]The Poisson pseudo-maximum likelihood (PPML) estimator accounts for heteroskedasticity bias and allows for trade flows to be zero. Silva and Tenreyro (2006) have shown that, under heteroskedasticity and due to Jensen's inequality, the use of the OLS estimator severely biases the coefficient on distance. Indeed, recent studies estimating the gravity equation rely almost exclusively on the PPML estimator, as it has been declared best practice in the gravity literature (Yotov et al., 2016).

version of equation (3.1). Equation (3.6) includes exporter-year $\gamma_{i,t}$ and importer-year $\delta_{j,t}$ fixed effects to account for income and expenditure, endogenous prices, and unobserved time-varying exporter and importer multilateral heterogeneity. By including a set of country-pair fixed effects ϕ_{ij} we control for unobserved heterogeneity across country pairs.[9]

The inclusion of time-invariant country-pair fixed effects captures all time-invariant factors, which means that we cannot estimate the distance elasticity and the border effect. However, we can observe time-varying changes in these bilateral trade costs by interacting $Dist_{ij}$ and $INTL\_BRDR_{ij}$ with a year dummy. The specification in equation (3.6) then allows for different effects of distance and border in each year $T \in \{1928, 1931, 1934, 1937\}$. These variables will capture all bilateral factors that depend on distance and borders influencing trade relative to the base year 1925.

Moreover, I include intranational trade and the variable $INTL\_BRDR_{ij,T}$, which indicates international trade, providing us with an estimate of the border effect and also ensuring that we measure international relative to intranational trade costs. This dummy variable, which takes the value of one for international trade ($i \neq j$) and zero for intranational trade ($i = j$), accounts for average increases across countries in unobservable export costs that decrease international trade relative to intranational trade. We expect the coefficient of this variable to be negative and increasing over time, capturing the effects of increasing capital controls, tariffs and other non-tariff trade barriers.

Finally, the variable *Cbloc* indicates whether two countries are on the gold standard or members of one of following trade or currency blocs: the sterling bloc, gold bloc, U.S. dollar bloc, Reichsmark bloc, exchange-control bloc, reciprocal trade agreements act (RTAA) or the IPS.

Indeed, the principle reason to use a panel approach is to get unbiased estimates for the gold standard and the trade and currency blocs. At least since Baier and Bergstrand's (2007) criticism, authors have been including country-pair fixed effects that control for potential endogeneity of trade agreements, currency areas or any form of economic integration area. All authors who investigated the interwar bloc system have confirmed the presence of strong endogeneity in these blocs. Eichengreen and Irwin (1995), Gowa and Hicks (2013) and Wolf and Ritschl (2011) all concur that the blocs are endogenous to preexisting trade flows among their members, reflecting rather than increasing their trade. However, equation (3.6) is an improvement over previous studies since it includes all features considered best practices in the gravity literature (Yotov et al., 2016).

First, I control for the border effect by including the variable $INTL\_BRDR_{ij,T}$. Bergstrand et al. (2015) have shown that the estimator of postwar currency unions is biased upward because it captures the average effects of globalization. Applied to the interwar period, the bloc dummy would be biased downward, capturing the average effect of deglobalization. Including the border dummy isolates the effect of trade and currency blocs on bilateral trade to determine how much a bloc increased trade between two members, but at the same time controls for increasing trends in unobservable bilateral trade costs that decreased international trade relative to intranational trade.

Second, I give trade flows the opportunity to adjust in a three-year interval. Trade policy changes will not be instantaneous and it is best practice among economists to give trade flows three to five years to adjust. Because the use of interval years comes at the cost of a decreased time variation, I

[9] Unless otherwise stated, I estimate equation (3.6) by using the fast PPML command provided by Larch et al. (2017) and limit pair fixed effects to be symmetric (i.e. $\phi_{ij} = \phi_{ji}$)

also estimate the model using data for all 13 years as a robustness check.

Third, as a further robustness check I allow for nonlinear effects of currency blocs to capture the possibility of the effects of blocs changing over time. This is done by including various lags ($Cbloc_{t-s}$) in the specification. There are two economic reasons why we should include lags. First, the bloc dummy is constructed using the dates of entry and exit. Academic debates over these dates put aside, it is reasonable to expect that the full economic effect of an exit from the gold standard or the entry into a new trade or currency bloc is felt only sometime after the event. Second, the collapse of the multilateral payment system into regional blocs alters the terms of trade and as is well known from the literature in international economics, terms-of-trade changes tend to have lagged effects on trade volumes.

Finally, I test whether the specification, through the inclusion of pair fixed effects, properly accounts for possible "reverse causality" between trade and bloc formation. I implement an easy test to assess the "strict exogeneity" of currency blocs by adding a new variable capturing the future level of currency blocs. A lead variable $Cbloc_{t+4}$ (4 years) of the bloc dummy is included in the specification to test for reverse causality. In the panel context here, if currency bloc changes are strictly exogenous to trade flow changes, $CBloc_{t+4}$ should be uncorrelated with the concurrent trade flow.

Table (1) presents the results from estimating different variants of equation (3.6). Column (1) shows the result for the main specification using the GE sample. The first thing to note is that the coefficient on distance is insignificant for all years except 1931. Because I control for any unobservables at the bilateral level by including pair fixed effects, this means that the distance coefficients describe the change in the distance elasticity relative to 1925.[10] Since the distance elasticity is negative, a negative coefficient in table (1) implies an increase (in absolute value) in the distance elasticity relative to 1925. The coefficient on $Dist_{ij,1931}$ in column (1) implies that the effect of distance had increased by 7.5% ($100 * (e^{0.0725} - 1$) in 1931.

The second finding is a very large and increasing border effect. The coefficient on $INTL\_BRDR_{ij,1937}$ implies that, all else being equal, the trade-decreasing effect of international borders had increased by 89.6% ($100 * (e^{0.64} - 1$) in 1937 relative to 1925.

The third finding in column (1) is that the coefficients on $Gold_{ij,t}$, $SterlingBloc_{ij,t}$ and $IPS_{ij,t}$ are large and significant. These coefficients state that being on gold increases trade between two members by 9.2% on average ($100 * (e^{0.088} - 1$), sterling bloc membership increases trade by 13% ($100 * (e^{0.122} - 1$) and IPS membership increases trade by 21% ($100 * (e^{0.191} - 1$). This is the baseline trade cost function that I use in the GE analysis in the next section. As discussed in section (3.1), the *GE sample* assumes a large number of missing observations to be zero. The remainder of this subsection therefore uses the partial sample, which lets us estimate the gravity model with more confidence.

Column (2) reestimates the main specification with the partial sample. The only difference in the estimated coefficients is the coefficient on sterling bloc membership, which is now insignificant.

[10] The pair fixed effects control for *initial* distance and border effects. As described in section 3.1, we assume that some factors do not change during this short time period. In that sense, the pair fixed effects control for much heterogeneity including nonlinearities in transport costs, impediments to information flows and home-variety biased preferences.

Table 1: Estimation of the Interwar Trade Cost Function

VARIABLES	(1) PPML X_{ij}	(2) PPML X_{ij}	(3) PPML X_{ij}	(4) PPML X_{ij}	(5) PPML X_{ij}
$\ln(Dist_{ij,1928})$	-0.0156	-0.0164	-0.0152	-0.0162	-0.0211
	(0.0275)	(0.0124)	(0.0124)	(0.0124)	(0.0132)
$\ln(Dist_{ij,1931})$	-0.0725**	-0.0734**	-0.0607**	-0.0619**	-0.0762***
	(0.0351)	(0.0288)	(0.0271)	(0.0270)	(0.0249)
$\ln(Dist_{ij,1934})$	0.0353	0.0328	0.0439	0.0701**	0.0630**
	(0.0284)	(0.0281)	(0.0281)	(0.0273)	(0.0280)
$\ln(Dist_{ij,1937})$	0.0600	0.0556	0.0671*	0.0947***	0.0915**
	(0.0368)	(0.0368)	(0.0369)	(0.0360)	(0.0373)
$INTL\_BRDR_{ij,1928}$	-0.0627	-0.0624*	-0.0625*	-0.0572	-0.00905
	(0.0612)	(0.0369)	(0.0370)	(0.0374)	(0.0335)
$INTL\_BRDR_{ij,1931}$	-0.289***	-0.289***	-0.302***	-0.296***	-0.232***
	(0.0839)	(0.0705)	(0.0686)	(0.0686)	(0.0585)
$INTL\_BRDR_{ij,1934}$	-0.669***	-0.671***	-0.680***	-0.709***	-0.718***
	(0.0635)	(0.0692)	(0.0703)	(0.0719)	(0.0723)
$INTL\_BRDR_{ij,1937}$	-0.640***	-0.633***	-0.644***	-0.675***	-0.698***
	(0.0897)	(0.0937)	(0.0947)	(0.0970)	(0.0988)
$Gold_{ij,t}$	0.0876**	0.0878**	0.0836**	0.0773**	
	(0.0397)	(0.0342)	(0.0339)	(0.0346)	
$SterlingBloc_{ij,t}$	0.122*	0.124			
	(0.0690)	(0.0761)			
$IPS_{ij,t}$	0.191**	0.194**	0.247***		
	(0.0826)	(0.0921)	(0.0846)		
Observations	6,090	5,085	5,085	5,085	5,085
Sample	*GE*	*Partial*	*Partial*	*Partial*	*Partial*
Country Pair Fixed Effects	Yes	Yes	Yes	Yes	Yes

Notes: All estimates are computed with data for the years 1925, 1928, 1931, 1934 and 1937, and use exporter-time, importer-time and pair fixed effects. Column (1) uses the PPML command, while all other estimations use the fast PPML command provided by Larch et al. (2017). All pair fixed effects are restricted to be symmetric (i.e. $\phi_{ij} = \phi_{ji}$). The estimates of fixed effects are omitted for brevity. Standard errors are clustered by country pair in parentheses; *** p<0.01, ** p<0.05, * p<0.1

This suggests that there is no market difference when using the more rigorous partial sample. The justification for including the two bloc variables $SterlingBloc_{ij,t}$ and $IPS_{ij,t}$ can be seen in columns (3) to (5). Excluding these two blocs in the specifications in column (4) and (5) changes the coefficient on distance, which now takes positive values for the years 1934 and 1937. For example, the coefficient on $Dist_{ij,1937}$ in column (4) suggests that the elasticity of trade to distance had declined by 9.9% ($100 * (e^{0.095} - 1)$ relative to 1925. This is a striking result since it means that distance was significantly less important as a trade cost in 1937 than in any other year in the sample, even though transport costs are said to have risen over the course of the interwar years. Importantly, this estimate is not biased by any other *average* increase in trade costs such as a general rise in tariff levels, which would render distance *relatively* less important. The decline in absolute value of the coefficient on distance has been observed by Eichengreen and Irwin (1995) and named the *interwar distance puzzle* by Albers (2018) since it resembles the post-war distance puzzle described by Disdier and Head (2008). Both Eichengreen and Irwin (1995) and Albers (2018) have argued that political trade barriers caused the decline in the distance elasticity. Contrary to these authors, I include intranational trade and a dummy for international borders, thereby measuring international relative to intranational trade costs while at the same time controlling for the average effect of deglobalization (i.e. the average increase in tariff and non-tariff trade barriers). This strategy has proved to be a solution to the post-war distance puzzle (Yotov, 2012; Yotov et al., 2016). In that sense, the results presented in columns (4) and (5) in table (1) should provide a solution to the interwar distance puzzle and the fact that the puzzle is not resolved indicates an omitted variable bias.

The puzzle is resolved once we add IPS and sterling bloc membership in columns (3) and (2). Controlling for these two blocs is important in order to get an unbiased estimate of the distance elasticity. Preferential tariff rates within the empire and fixed exchange rates between sterling bloc members rendered distance within these blocs (and consequently also between non-members) less important. I therefore include sterling bloc membership in the trade cost function although the coefficient on $SterlingBloc_{ij,t}$ in column (2) is just not significant at the 10% level.

This solution to the interwar distance puzzle is robust when extending the partial sample to include all 13 years, interacting the pair fixed effects with a time trend, using asymmetric pair fixed effects or using lagged variables for the bloc dummies (see appendix (A.2)). Because of the large number of fixed effects and variables to be estimated with a relatively small number of observations, compared to post-war trade studies, I also estimate the gravity model by excluding the pair-fixed effects and including traditional gravity covariates such as colonial ties (see A.3). Here again, we observe the disappearance of the distance elasticity puzzle once we include sterling bloc and IPS membership. Moreover, in many specifications in the robustness appendix, the coefficient on sterling bloc membership is significant at the 10% level. Overall, the results presented in table (1) and the choice of the trade cost function are robust to the various specifications.

The decision to include IPS and sterling bloc membership in the regressions in table (1) is based on the regression results in table (2), which show that these blocs are the only ones that are significant and not subject to reverse causality. Table (2) uses equation (3.6) to estimate the effects of the currency blocs. We drop $Dist_{ij,t}$ and $Gold_{ij,t}$ for simplicity, but include $INTL\_BRDR_{ij,t}$ to control for the average effect of deglobalization. Column (1) includes the trade blocs IPS, exchange control bloc, RTAA and the gold bloc, column (2) includes the currency blocs sterling bloc, Reichsmark bloc, gold bloc and dollar bloc, and column (3) includes all blocs. RTAA, gold bloc and dollar bloc are insignificant in all specifications and are excluded from further analysis.

Column (4) tests the remaining four blocs for reverse causality by including a four-year lead variable for each bloc. These lead variables should be insignificant in the absence of any reverse causality. The test suggests that only the IPS and the sterling bloc are free from potential endogeneity issues. Although the sterling bloc dummy is insignificant in columns (3) and (4), I include it in the trade cost function for two reasons. First, adding the gold standard dummy, lags of the bloc variables, and distance drastically improves the significance of the sterling bloc (see the appendix (A.2)). Second, including sterling bloc membership in the trade cost function solves the distance puzzle as we saw in table (1).

Overall, the results suggest a strong trade-increasing effect of the IPS that is economically and statistically significant in all specifications. This stands in contrast to the results obtained by Gowa and Hicks (2013) who find that not a single bloc increased member trade. Instead, my results support De Bromhead et al. (2019) recent finding that the IPS was successful in shifting trade towards the empire. More importantly, the results warrant the use of the variable in the regressions presented in table (1). To answer the question of how much the collapse of the gold standard contributed to the collapse of world trade in the 1930s, one needs to consider deducting the trade-increasing effect of these blocs, since without the collapse of the gold standard, the IPS and the sterling bloc might not have formed.

Table 2: Trade and Currency Blocs in the Interwar Period

VARIABLES	(1) PPML X_{ij}	(2) PPML X_{ij}	(3) PPML X_{ij}	(4) PPML X_{ij}
$INTL\_BRDR_{ij,1928}$	-0.0528***	-0.0541***	-0.0539***	-0.0793***
	(0.0155)	(0.0154)	(0.0154)	(0.0195)
$INTL\_BRDR_{ij,1931}$	-0.390***	-0.397***	-0.404***	-0.430***
	(0.0282)	(0.0294)	(0.0284)	(0.0309)
$INTL\_BRDR_{ij,1934}$	-0.630***	-0.549***	-0.570***	-0.581***
	(0.0362)	(0.0435)	(0.0437)	(0.0441)
$INTL\_BRDR_{ij,1937}$	-0.556***	-0.467***	-0.491***	-0.504***
	(0.0420)	(0.0433)	(0.0444)	(0.0442)
$SterlingBloc_{ij,t}$		0.133**	0.0692	0.102
		(0.0654)	(0.0674)	(0.0756)
$IPS_{ij,t}$	0.331***		0.228***	0.212***
	(0.0755)		(0.0755)	(0.0735)
$ExchangeControlBloc_{ij,t}$	0.0643		0.265***	0.318***
	(0.0740)		(0.0797)	(0.0808)
$RMBloc_{ij,t}$		-0.516***	-0.566***	-0.685***
		(0.131)	(0.130)	(0.123)
$SterlingBloc_{ij,t+4}$				-0.0401
				(0.0686)
$IPS_{ij,t+4}$				-0.0303
				(0.0385)
$ExchangeControlBloc_{ij,t+4}$				-0.223***
				(0.0540)
$RMBloc_{ij,t+4}$				0.362***
				(0.0823)
$RTAA_{ij,t+4}$	0.0926		0.0339	
	(0.0818)		(0.0852)	
$GoldBloc_{ij,t}$	0.0129	-0.000762	0.00619	
	(0.0368)	(0.0377)	(0.0378)	
$DollarBloc_{ij,t}$		0.00213	0.00232	
		(0.0847)	(0.0945)	
Observations	5,085	5,085	5,085	5,084
Country Pair Fixed Effects	Yes	Yes	Yes	Yes

Notes: All estimates are computed using the partial sample and the years 1925, 1928, 1931, 1934 and 1937. All regressions include exporter-time, importer-time and pair fixed effects. The estimates of fixed effects are omitted for brevity. All regressions are estimated using the fast PPML command provided by Larch et al. (2017) and pair fixed effects are restricted to be symmetric (i.e. $\phi_{ij} = \phi_{ji}$). Standard errors, in parentheses, are clustered by country pair; *** p<0.01, ** p<0.05, * p<0.1.

3.3 General Equilibrium Analysis

The analysis so far has focused solely on partial effects of distance, borders and the payment system. Using the gravity model in equations (3.1) to (3.5) I now undertake a quantitative general equilibrium (GE) comparative static exercise to get a complete evaluation of the impact of changing trade costs on the interwar trade bust.

As discussed in the previous section, a large part of trade data is missing or equal to zero for a given pair over the whole period of investigation. This makes it impossible to identify and obtain the estimates of the complete set of pair fixed effects, which are used to construct bilateral trade costs. I deal with the issue by adopting the two-stage procedure proposed by Anderson and Yotov (2016). First, I estimate my preferred specification of the gravity model in order to obtain the estimates of the bilateral fixed effects (ϕ_{ij}) for country pairs with at least one non-zero trade flow. The estimation results for this regression are shown in column (1) of table (1). For these observations, the following trade cost function is calculated for the year 1937 as the baseline scenario:

$$\begin{aligned} \tau^{BLN}_{ij,1937} = exp(\hat{\phi}_{ij} + 0.06 \ln Dist_{ij,1937} - 0.64 * INTL\_BRDR_{ij,1937}) * \\ exp(0.088 * Gold_{ij,t} + 0.191 * IPS_{ij,t} + 0.122 * SterlingBloc_{ij,t}) \end{aligned} \tag{3.7}$$

In the second step, I regress the estimates of pair fixed effects on distance, contiguity, colonial linkage, common language, the border dummy, and exporter and importer fixed effects:

$$\begin{aligned} \hat{\phi}_{ij} = exp[\alpha_1 * \ln Dist_{ij} + \alpha_2 * Contig + \alpha_3 * Colonial + \alpha_4 * ComLang] * \\ exp[\alpha_4 * INTL\_BRDR_{ij} + \gamma_i + \delta_j] + \epsilon_{ij} \end{aligned} \tag{3.8}$$

The predicted pair fixed effects from this second stage regression are used to fill up the missing pair fixed effects in order to construct the complete set of bilateral trade costs that can then be used as the baseline scenario in the counterfactual analyses. I now reestimate the gravity model constrained with the complete set of bilateral trade costs for the year 1937. The estimates of the exporter and importer fixed effects from this regression are used in order to construct all baseline values of the inward and outward multilateral resistance terms, which in turn are used in combination with data on output and expenditure to obtain the GE indices of country i's overall trade in the baseline.

Next, we define three counterfactual scenarios (CF), which translate into three counterfactual trade cost functions. The first CF assumes that the gold standard, in its 1928 form, still existed in 1937 and that the IPS and sterling bloc had not formed. The conditional GE is achieved by reestimating the econometric gravity specification for the year 1937 under the following constraint

$$\begin{aligned} X_{ij} = exp[0.06 \ln Dist_{ij,1937} + -0.64 * INTL\_BRDR_{ij,1937}] * \\ exp[0.088 * Gold^{CF}_{ij,1928} * + \gamma^{CF}_i + \delta^{CF}_j + \hat{\phi}_{ij}] + \epsilon^{CF}_{ij} \end{aligned} \tag{3.9}$$

where the bloc variables now take their 1928 values. The CF then describes a scenario where gold standard adherence had remained at its 1928 level and the IPS and sterling bloc had not formed. The predicted volume of trade from regression (3.9) is used to calculate country i's counterfactual conditional GE trade volume ($\hat{X}_i^{CF} = \sum_{j=1}^{N} \hat{X}_{ij}^{CF}$ for all $j \neq i$). The new set of estimates of exporter and importer fixed effects from specification (3.9) and the constrained coefficients of the trade cost variables are used to construct the corresponding conditional GE multilateral resistances and obtain real GDP estimates for each of the 36 countries in the sample.[11] Finally, the effects of the full endowment GE are obtained by implementing a four-stage iterative procedure that allows for endogenous factory-gate prices, income, expenditure and trade to adjust to the counterfactual shock. The value for the elasticity of substitution is 7, which we take from the literature.[12] A detailed description of the calculation of the GE effects is summarized in appendix (A.3).

The second CF assumes that the trade-decreasing effect of borders had not changed in 1937 relative to 1928. Hence, we constrain the coefficient on $INTL\_BRDR_{ij,1937}$ to take its 1928 value (-0.063). The third CF constrains the coefficient on $\ln Dist_{ij,1937}$ to be equal to -0.016, the coefficient for 1928. Table (3) presents the results and shows how each country's total trade (in 1937 prices) would have changed under each of the three CF scenarios. The two lines at the bottom of the table show average change across countries and change in total trade aggregated over all 36 countries, which represent the counterfactual change in world trade.

The results are striking and show a clear winner of the horse race. Had the average level of tariff and non-tariff trade barriers, as captured by the border effect, not changed relative to 1928, world trade would have been 64.6 % larger. Crossing the finish line second is the collapse of the gold standard. Had countries remained on the common payment system and had the sterling bloc and IPS not formed, trade would have been 3 % larger. Finally, transport costs, as captured by distance, is not last but has to be disqualified as a cause of the great trade collapse since it ran off in the opposite direction. Had the distance elasticity remained the same, total trade would have been 19% lower on average. Since the baseline and counterfactual trade costs in this last CF are calculated with insignificant coefficients on the distance elasticty (see section (3.2)), the GE effects in CF 3 are likely to be zero.

There is a significant amount of variation of the impact these CF scenarios would have had across countries. A lower border effect would have benefited South Africa most, roughly doubling her trade. Great Britain, Italy and France would also have seen above average trade increases. Germany and the U.S., on the other hand, would have seen relatively small increases in their aggregate trade had borders remained at their 1928 level. The gold bloc countries (France, Belgium, Netherlands, Switzerland and Italy) and most of Latin America would have particularly benefited, had the gold standard not collapsed. Finally, we observe that, with the exception of Canada, the trade-increasing GE effect in CF 1 is negative for Great Britain and its colonies and dominions. This means that the trade-increasing effect of the IPS and sterling bloc outweighed the negative impact of the collapse of the gold standard for those countries.

[11] I only present total exports. Other indices are available upon request.

[12] The criteria of convergence are set so that either the standard errors or maximum of the difference between two iterations of the factory-gate prices are smaller than 0.01. All three scenarios in table (3) have also been calculated with an elasticity of substitution of 5, another common value in the literature. The results did not significantly change.

Table 3: General Equilibrium Comparative Statics

Country	CF 1: Gold		CF 2: Border		CF 3: Distance†	
	CDL GE	Full GE	CDL GE	Full GE	CDL GE	Full GE
ARG	2,49	2,52	46,51	48,94	-13,53	-20,83
AUS	-6,06	-6,79	50,28	51,96	-14,42	-21,83
AUT	8,02	7,97	71,15	71,17	-10,27	-16,08
BEL	7,87	7,86	61,47	61,94	-11,32	-16,58
BGR	4,37	4,82	26,30	31,07	-7,44	-14,20
BRA	8,93	8,81	73,90	73,31	-15,89	-22,44
CAN	1,64	1,48	54,45	56,67	-6,46	-14,32
CHE	9,09	8,86	74,89	73,52	-12,32	-17,51
CHL	4,01	4,69	35,39	40,17	-9,21	-17,64
COL	9,09	9,03	75,26	75,14	-16,19	-22,09
DEU	6,94	7,06	57,94	59,04	-8,83	-15,26
DNK	2,01	1,86	64,42	64,30	-10,43	-15,99
EGY	2,47	2,34	59,56	59,83	-16,40	-22,00
ESP	-0,29	-0,38	80,60	77,79	-12,81	-18,65
FIN	2,26	2,20	53,26	54,76	-11,08	-17,30
FRA	9,86	9,46	82,34	79,33	-11,71	-17,55
GBR	-3,93	-4,02	80,53	78,36	-18,10	-23,15
GRC	9,61	9,33	78,97	77,09	-18,52	-23,26
GTM	8,25	8,35	68,20	69,49	-20,64	-25,57
HND	-0,05	-0,02	65,09	66,52	-15,47	-21,22
HUN	5,44	5,85	48,38	51,79	-6,57	-13,07
IDN	5,00	5,25	46,42	49,40	-12,73	-20,25
IND	-4,72	-5,28	55,95	57,09	-13,42	-20,51
ITA	9,97	9,57	84,26	81,44	-13,50	-19,25
JPN	-0,03	-0,12	76,35	75,22	-17,41	-23,02
KOR	-0,05	-0,14	73,85	72,31	-13,78	-19,00
MEX	9,66	9,45	82,01	80,58	-12,87	-19,31
NLD	8,49	8,46	66,62	66,91	-11,27	-16,57
NOR	2,84	2,59	81,62	78,70	-12,58	-18,14
NZL	-11,73	-12,51	50,68	51,69	-18,91	-25,28
PRT	-5,48	-5,54	78,74	75,66	-16,04	-21,12
SWE	3,10	2,96	64,53	64,60	-10,30	-16,42
URY	6,96	7,14	57,77	59,05	-18,46	-24,44
USA	6,39	6,53	57,55	59,18	-10,27	-18,09
YUG	-0,28	-0,41	48,56	51,43	-7,49	-14,00
ZAF	-13,01	-12,66	101,76	94,01	-27,95	-31,80
Country Average	3,03	2,96	64,88	64,99	-13,46	-19,55
Aggregate Change	3,03	2,95	64,39	64,58	-12,59	-18,98

Note: This table reports the GE trade effects of changing three components of the trade cost function in equation (3.7). The first scenario (CF 1) assumes that the gold standard had not collapsed and the IPS and sterling bloc had not formed. The second scenario (CF 2) assumes that the border effect had remained at its 1928 value, and the third scenario (CF 3) assumes that the distance coefficient had remained at its 1928 value. For all three scenarios, I report two different trade impacts: the conditional GE trade impact (CDL GE), which takes changes in the multilateral resistances (MR) into account, but holds GDPs constant; and the full endowment GE trade impact (Full GE), where MRs and GDPs adjust. The row "Country Average" shows the arithmetic average of the GE effects of all countries and the bottom line "Aggregate Change" states the total impact on aggregated trade.
† CF 3 is computed using insignificant coefficients. Hence, the GE effects are also insignificant.

These results are robust to the use of different parameters and counterfactual scenarios. I performed the same GE analysis above using column (3) in table (1) as the trade cost function, setting sigma equal to 5 and assuming as a CF that the IPS had formed even if the gold standard had not collapsed. [13]

[13] For brevity, these results are not shown here, but are available upon request. If anything, these alternative specifications provide even stronger evidence against transport costs as a cause of the low levels of world trade in 1937.

4 A New Transport Cost Index

Section (3) established that the distance elasticity increased (in absolute value) in 1931. However, after 1931 the distance elasticity decreased, so that in 1937 distance did not matter any more than it did in 1925. If we interpret the distance elasticity as a proxy for transport costs, the estimates suggest that transport costs rose during the worst years of the Great Depression but then returned to their pre-depression level. This section compares this result with evidence on real freight rates.

There has been considerable disagreement about the course of transportation costs in the interwar period (Estevadeordal et al., 2003; Mohammed and Williamson, 2004; Hynes et al., 2012; Albers, 2018). Much of the discussion about transport costs revolves around the Isserlis Index, the standard source of global freight trends, and the choice of the price deflator. Depending on the frequency of the data and the choice of the deflator, the Isserlis Index implies rising or falling transport costs. For example, Estevadeordal et al. (2003), who use the Sauerbeck consumer price index to deflate the Isserlis index, find rising real freight rates for the whole of the interwar period. However, the use of the Sauerbeck index is problematic since it includes non-tradable goods. Mohammed and Williamson (2004) make use of the original source of the Isserlis index, Angier's annual reports on British shipping, and construct route-specific deflators. Relative to the 1920s, their real freight rate index shows a fall between 1930 and 1934 and a rise between 1935 and 1939. Unfortunately, they only provide five-year averages and we cannot say how much of the increase during the late 1930s was driven by the year 1939, the start of World War II. A more general caveat concerning the use of the Angier data is that one relies on freight rates for British tramp shipping to make inferences about the general evolution of transport costs. Their index ignores the liner shipping industry, which carried high value articles, whereas tramps carried the high bulk, low value staples. Furthermore, liners, contrary to tramps, operated on fixed routes and fixed schedules. Therefore, we cannot assume that Mohammed and Williamson's (2004) index is representative for the cost of shipping high value manufactures or for the entire British shipping industry. It also ignores transportation industries that transported goods on railroads, turnpikes, rivers and planes. Moreover, the Mohammed and Williamson (2004) index relies heavily on routes to and from Britain and completely ignores shipping between non-European ports.

Albers (2018) presents new freight data for wheat along four oceanic routes deflated by the price of the good at the place of origin. Additionally, he presents data on German railway freight rates deflated by the German wholesale price index. His series imply a modest but economically significant increase in real transport costs over the period from 1925 to 1936 and a spike around 1931. This mixed evidence calls for further evidence on the development of transport costs in the interwar period, which I provide with a new index on freight rates for US cotton.

I compiled monthly data on cotton (American middling) freight rates from New York to 21 destinations (for high- and low-density cotton) from the Commercial and Financial Chronicle (1925 - 1936). The data was published at least twice a month and I collected the data that is closest to the middle of the month.[14]

I deflate the arithmetic average between high- and low-density nominal freight rates (ct per lb)

[14] Six routes are not covered for the whole period and either enter the series in 1926 (Venice) or stop reporting sometime in the 1930s (Lisbon, Oporto, Barcelona, Japan, Shanghai).

by the monthly price (ct per lb) of middling upland in New York, which I take from the Statistisches Reichsamt (1936), to get freight rates in ad-valorem terms. Figure (A.2) in the appendix plots these ad-valorem freight rates for all 21 routes. The graphs show that real transport costs of cotton did indeed rise during the Great Depression. At the beginning of 1925 ad-valorem rates ranged from 1.5% for Barcelona to 3.3% for Salonica. In 1928 ad-valorem rates ranged from 1.7% to 5% and in June 1932, at the height of the depression, rates had increased to 6.5% in Le Havre and 15.6% in Salonica. However, by December 1936 most European rates had fallen to 3% again, while Salonica stayed at 7.2%. Between 1928 and 1936 average ad-valorem freight rates increased by only 1.2 percentage points. There is a strong positive relationship between these route specific freight rates and distance. A simple OLS regression suggests that the elasticity of ad-valorem freight rates to distance is between 0.4 and 0.5 (see table (A.6) in the appendix).

To examine how much transport costs rose when expressed as an index, I weight these routes by their export quantities to create a Laspeyres-type index:

$$F_t = \frac{\sum (f_{n,t} * q_{n,t})}{\sum (f_{n,10.1927} * q_{n,10.1927})} \tag{4.1}$$

where $f_{n,t}$ is the ad-valorem freight rate from New York to location n in month t and $q_{n,t}$ is the quantity (as a percentage of the total, i.e. $\sum_n q_{n,t} = 1$) exported to location n in month t. Quantities come from the Bureau of Foreign and Domestic Commerce (October, 1927) and I use October 1927 as the base month.[15] Figure (2) below depicts the freight index from 1925 to 1936.

The first thing that deserves mention is the spike in the fall of 1926. The index increases by 100 % between mid-August and mid-November. This increase in real freight rates comes at the end of the British coal strike of 1926, which had started in May that year. After the strike ends, the index falls again and remains relatively stable until the summer of 1930. The index reaches its peak in June 1932 at the height of the worldwide deflation and falls sharply afterwards. In 1936 the index is on average 20 percentage points above the base period (October 1927) and 50 percentage points higher than the index average in 1925. However, this change is only large if expressed as an index. As we have seen above, freight rates in ad-valorem terms increased by less than two percentage points.

A difference of this index when compared with Mohammed and Williamson (2004) is that it relies heavily on the liner industry and is based on routes from New York to cities in Europe and Asia.[16] The index therefore should serve as a useful supplement to the British tramp shipping index of Mohammed and Williamson (2004). The drawback of this index is that its construction is based on only one commodity. It is thus difficult to make judgments about Estevadeordal et al.'s (2003) different hypotheses on the causes of the rise in shipping costs. On the one hand, real freight costs rose slightly over the entire period suggesting that productivity growth in the shipping sector was slower than productivity growth in the cotton sector. On the other hand, estimating the elasticity

[15] I use exports of unmanufactured cotton and match the export destination country with the destination port in the freight rate data. Whenever more than one route goes to the same country of destination, I allocate an equal share of the quantity to each port (e.g., I assume that 50% of cotton exports to Germany arrive in Hamburg and 50% in Bremen). For Fiume, Piraeus and Salonica I use exports to "other European countries".

[16] Liners that operated frequently on these routes include RMS Ascania, RMS Scythia and RMS Aurania of the Cunard line (see the Commercial and Financial Chronicle (1925 - 1936)).

Figure 2: Cotton Freight Index

Notes: The graph shows an index of cotton freight rates calculated using the Laspeyres Index for 21 ocean routes. Data on freights rates, prices and quantities come from the Commercial and Financial Chronicle (1925 - 1936), the Statistisches Reichsamt (1936) and the Bureau of Foreign and Domestic Commerce (October, 1927).

of freight rate to distance, a proxy for technology in the shipping industry, shows that this elasticity is lower in the post- than in the pre-depression period (see appendix (A.4)).

Their alternative explanation for rising shipping costs, rigidities that prevented nominal freight rates to adjust (e.g. shipping cartels), is clearly visible in figure (2). Overall, the index in figure (2) and the ad-valorem freight rates along different routes in figure (A.2) are broadly in line with the data on transport costs by Albers (2018). Furthermore, the large increase in real transport costs during the extreme deflation, between 1929 and 1933, resembles the increase in the distance elasticity in 1931 (see section (3)).

How does the evidence on transport costs compare with the regression results in section (3.2)? The assembled data in this section suggests that transport costs between 1933 and 1936 were not significantly different from those in the 1920s, at least in any economic sense. Although they may have contributed to the initial trade bust during the first years of the depression, the marginally higher level of transport costs, relative to the 1920s, cannot explain the low levels of international trade still present in the latter half of the 1930s. This is the same result I obtained econometrically in columns (1) and (2) of table (1), where the change in the distance elasticity in 1937 relative to 1925 is insignificant.

Although insignificant for the preferred specification in table (1) column (2), the coefficient on $Dist_{ij,T}$ is actually positive, suggesting that the distance elasticity had decreased by 5.7% ($100 * (e^{0.0556} - 1)$ relative to 1925. This effect becomes more significant if we use 1928 as the base year or drop the sterling bloc dummy. How does this more critical interpretation of the regression results square with the fact that we actually observe a small increase in transport costs?

A sensible explanation for this is that freight rates themselves do not perfectly describe transportation costs. Transportation of goods involves time, which increases with distance between locations. Hummels and Schaur (2013) have argued that time in transit is equivalent to an ad-valorem tariff of 0.6 to 2.1%. Although time as a trade barrier is likely to be more important in the age of global supply chains, one cannot disregard technological improvements in the transportation industry, the building of highways and the advent of aviation during the interwar period. Indeed, the regression results in table (A.6) suggest an improvement in the technological relationship between freight rates and distance over the interwar period.

A second reason why freight rates might not adequately describe the real costs of transportation is that during the depression, many governments started to subsidize freight costs. In 1933, the U.S. Secretary of Agriculture (1933) complained to Congress:

> Indirect export subsidies are sometimes granted by governments that operate the railways in their territory in the form of specially reduced freight rates. Reduced rates for export shipments apply, for instance, to wheat in India, sugar in Germany, corn in Rumania, and hops in Czechoslovakia.

Naturally, if governments subsidized freight costs, the data on freight rates will not be what exporters pay to ship their goods. Instead, increasing subsidies on the transportation of goods would render distance less important as a trade cost, which is precisely what a more critical interpretation of the regression results in section (3.2) suggests.

5 Conclusion

This paper revisits the debate about the causes of the collapse of world trade in the 1930s and examines the relative importance of higher transport costs, the collapse of the payment system, and increased tariff and non-tariff trade barriers. Using a fully specified gravity equation motivated by formal theoretical foundations, I estimate the effects of bilateral distance, international borders, and the payment system on trade.

I show that the gold standard increased trade among its members *only* by 9.2%. The negative effect that the collapse of the gold standard into a system of trade and currency blocs had on world trade is further reduced when we take into account the trade-increasing effects of the IPS and the sterling bloc. Distance, a proxy for transport costs, did not matter more in 1937 than it did in 1925 and thereby fails to explain the low levels of world trade in the 1930s. The factor that changed dramatically from 1930 onwards is the border effect. The trade-reducing effect of international borders captures all trade costs, unrelated to distance (transport costs) or network effects (payment systems). The border effect then is the combined effect of all commercial and financial policies (e.g. tariffs, import quotas, capital controls). This already large trade reducing effect, increased by 89% from 1925 to 1937. Had countries not resorted to beggar-thy-neighbor policies and tariff retaliation after 1928, world trade would have been 64% larger.

The result that transport costs did not matter for the low levels of world trade in the late 1930s is supported by new data on ad-valorem freight rates. Shipping cotton on ocean liners was only slightly more expensive in the late 1930s than in the 1920s and it is unlikely that this marginal increase was the result of slow productivity growth. Transport costs and the distance elasticity, however, show strong increases around 1931, a period of severe deflation, which suggests that cartelization in the shipping industry did in fact matter. In that sense, transport costs might have mattered for the initial trade collapse at the beginning of the Great Depression. Future research would do well, in estimating a dynamic macroeconomic model for the years 1929 to 1933 to evaluate the relative importance of commercial policies, transport costs and credit frictions during these years.

This study also provides a novel contribution to the quest for an unbiased estimate of the distance elasticity in gravity models. I argue that the retreat into the British Empire decoupled a large part of the world's multilateral trade and payment system. Overall, this made distance less important, relative to other trade costs. Only after I control for the gold standard, the IPS and the sterling bloc does the evolution of the distance elasticity approximate the evolution of transport costs in the interwar period. Future work should test whether harmonization of political trade barriers add to the solution of the post-war distance puzzle.

A Appendix

A.1 Gravity Data Appendix

I created a dataset by merging two existing datasets. I use the large dataset by Fouquin and Hugot (2016b) and merge it with the dataset provided by Gowa and Hicks (2013).[17] Because I needed observations for internal trade, I restricted the sample to include countries with observations on GDP.[18] The 36 countries included and their bloc membership are listed in table (A.1).

Table A.1: Countries in both samples and bloc membership

Country	Bloc	Country	Bloc
Argentina	St (1934)	India	St (1931), IPS (1932)
Australia	St (1931), IPS (1932)	Indonesia	
Austria	Ex (1931), RM (1932)	Italy	G (1931-34), Ex (1931)
Belgium	G (1931-35)	Japan	
Bulgaria	Ex (1931), RM (1932)	Korea	
Brazil		Mexico	
Canada	IPS (1932)	Netherlands	G (1931-36)
Chile		Norway	St (1933)
Colombia		New Zealand	St (1931), IPS (1932)
Denmark	St (1933)	Portugal	St (1931)
Egypt	St (1931)	South Africa	St (1931), IPS (1932)
Finland	St (1933)	Spain	
France	St (1938), G (1931-36)	Sweden	St (1933)
Germany	Ex (1931), RM (1932)	Switzerland	G (1931-36)
Greece	Ex (1931), RM (1932)	United Kingdom	St (1931), IPS (1932)
Guatemala		United States	
Honduras		Uruguay	
Hungary	Ex (1931), RM (1932)	Yugoslavia	Ex (1931)

Notes: St: Sterling bloc; G: Gold bloc; Ex: Exchange bloc; RM: Reichsmark bloc; IPS: Imperial Preference System. Data on bloc membership is taken from Gowa and Hicks (2013). Dollar bloc and RTAA membership is not reported here for brevity.

Next, I constructed two samples, one for the estimation of the partial effects in section (3.2) and one for the GE analysis in section (3.3), which I call *the partial sample* and *the GE sample*. The partial sample balances the panel, which means that for a given dyad an observation exists for every year from 1925 to 1937. Because the Fouquin and Hugot (2016b) dataset sometimes also reports observations with missing trade flows, *the partial sample* drops the country pairs that report missing trade flows for all years in the sample. I replaced the remaining 16 observations missing trade flows with the variable $FLOW_0$, which is equal to zero when it is reasonable to assume that the trade flow is missing because the trade flow is actually zero.[19]

[17] I used the yearly British pounds per dollar exchange rate in Fouquin and Hugot (2016b) to convert the Gowa and Hicks (2013) trade flows into pound sterling.

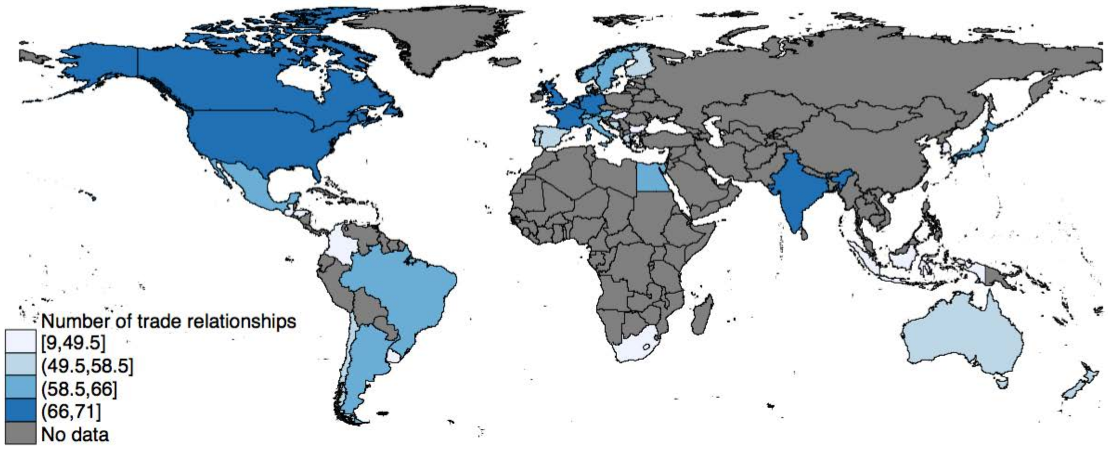
[18] Despite existing data on GDP, I excluded the only communist country, the USSR, from the sample.

[19] For details on the construction of the variable $FLOW_0$, see Fouquin and Hugot (2016b). The 16 observations that were replaced with zero are: IND - HND (1925, 1928); PRT - FIN (1937); BGR - IND (1925, 1928, 1937); KOR - IND (1928, 1931, 1934, 1937); GTM - IND (1931, 1934, 1937); COL - IND (1937); HND - NLD (1925); GTM - NLD (1925).

Figure (A.1) shows the geographical distribution of the 1017 dyads in *the partial sample*. The sample covers most of Europe and the Americas.

Figure A.1: Geographical distribution of *the partial sample*

[Print in Color]



Notes: This figure plots the geographical distribution of the sample on a map with current frontiers. The colors indicate the number of times a country appears in the observations. The partial sample includes 1017 country pairs from a total of 36 countries over five time periods. The map shows the number of trade relationships of a country in the sample. For example, Korea appears only in 9 dyads and is depicted in white.

The GE sample is obtained not by dropping dyads with missing observations in a given year, but instead by adding observations to obtain a quadratic matrix of $N * N$ observations per year (i.e. a quadratic matrix of trade relations between the N trading partners). This consequently results in 1296 observations per year or 6480 observations for 5 years, of which 881 missing observations are assumed to be zero.[20] These 881 missing observations are primarily country pairs of two distant small developing countries (e.g. Korea and Honduras) and it is reasonable to assume that these trade flows are zero or close to zero.

While it is possible to restrict the number of countries to reduce the number of missing observations, this comes at the cost of biasing the sample towards industrialized nations and reducing the total number of observations, which creates convergence problems. Note that biasing the sample heavily towards European countries increases the share of manufactures in aggregate trade as can be seen in table (A.2), which is taken from Hilgerdt (1942). Manufactured goods made up a larger share of trade (exports plus imports) of European countries compared to the rest of the world before and after the Great Depression. This affects the estimation of the distance elasticity through two channels, the elasticity of substitution and the elasticity of trade costs with respect to distance.[21]

[20] Note that the partial sample contains only 5085 observations. This is because the partial sample loses 514 non-missing observations in the process of balancing the panel.

[21] In a previous version of this paper I estimated the gravity equation in yearly cross-sections for different samples. One observes that the "distance puzzle" is more pronounced if the sample is restricted to include primarily European countries (results are available upon request). This has also been observed by Albers (2018). This result has its analogy in the literature on the post-war distance puzzle. In a recent study, Borchert and Yotov (2017) show that on average the distance elasticity has fallen between 1986 and 2006, but that low-income countries have not seen a fall in the distance elasticity. The authors argue that the distance elasticity depends heavily on the composition of exports,

Table A.2: Percentage Composition of Merchandise Trade by Groups of Countries

	Class	Imports 1928	Imports 1935	Exports 1928	Exports 1935
Europe	c	32.00	31.33	50.67	50.67
	a+b	68.00	68.67	49.33	49.33
Other countries	c	57.50	57.14	15.00	16.21
	a+b	42.14	42.86	85.00	83.79

Notes: a: Foodstuffs and live animals; b: Materials raw or partly manufactured; c: Manufactured articles; *Source*: (Hilgerdt, 1942, Table 7 p.23)

σ is different for the two classes of goods, but the importance of transport costs also differs between the two classes of goods. Freight rates are usually measured per unit of the good in question and depend on how bulky the good is. Arguably, the value of a cargo full of manufacture goods should on average be much higher compared to the value of the same cargo full of primary products. The share of shipping costs in the price of manufactures should then be much lower than in primary goods. The elasticity of trade in agricultural goods to transport costs is therefore larger compared to manufactured goods. It might be that the increase in real freight costs during the depression hit the trade of primary goods particularly hard.

Moreover, deflation in agricultural and primary products was more severe than in the manufacturing sector. Between 1925 and 1936, the export price ratio of manufactures to other goods increased by 29% (League of Nations, 1938).[22] Transport costs increased because real freight costs increased. It therefore matters that the deflation was more severe in the primary sector.[23] Dropping the countries with fewer observations (Korea, Honduras, Guatemala and Colombia) might therefore bias the results towards the "tariff explanation".

Both samples include observations for intranational trade (i.e. how much a country trades within its own borders). Lacking data on interregional trade or a measure of gross output (subtracting total exports from gross output would yield intranational trade), one needs another way of constructing a proxy for internal trade.[24] Jacks et al. (2011) simply use the GDP series as a proxy for gross output, which poses two problems. First, gross output is by construction larger than value-added GDP and so the use of GDP would lead to an underestimation of domestic trade. Second, as GDP includes services, which are not covered by the trade data, this leads to an overestimation of domestic trade. Fouquin and Hugot (2016a) follow a different approach. They scale up their GDP series by a factor of 3.16, which is the average ratio of gross output to value added from a post-1980 dataset. Here, I propose a more reasonable scale factor with which to multiply the GDP series. I use data from the US United States Department of Commerce ((1935 - 1937) of the interwar period and data from

in particular the value-to-weight ratio.

22 This is based on unit values; see (League of Nations, 1938, Table 1).

23 Deflation also increases the effective tariff rate. Whether real freight rates increase more than the effective tariff rate depends on the stickiness of nominal freight rates and the rate at which tariffs increase in the respective sector.

24 Indeed Bulgaria's GDP is smaller than her total exports for all years in the sample, which means that using GDP as a proxy of gross output would result in negative internal trade.

Federico (2004) to calculate average gross output to value added ratios for the U.S. manufacturing sector and international agricultural sector, and calculate internal trade as follows:[25]

$$X_{iit} = Y_{it}[(1 - s_{1,t} - s_{2,t}) + (s_{1,t} * v_{Agri,t}) + (s_{2,t} * v_{Manuf,t})] - X_{it} \tag{A.1}$$

In the above equation, Y_{it} is country i's GDP in year t. $s_{1,t}$ and $s_{2,t}$ are the global average shares of the primary and secondary sectors, which I take from Fouquin and Hugot (2016b). $v_{Agri,t}$ and $v_{Manuf,t}$ are the average ratios of gross output over value added. On average this yields a scale factor of 1.4. X_{it} is total exports and is taken from Fouquin and Hugot (2016b) for all countries but Yugoslavia. Total exports for Yugoslavia come from (Mitchell, 1998, p.580).[26]

In sections (3.2) and (3.3) I include a set of dummy variables that indicate whether two countries are on the gold standard or are members of the same trade or currency bloc. Data for the time on gold comes from Eichengreen (1992) and bloc membership is taken from Gowa and Hicks (2013).

Finally, I include a set of standard gravity covariates (distance, contiguity, common language, colonial linkage). I use the existing data on these variables in Fouquin and Hugot's (2016b) dataset, which I supplement with the CEPII distance dataset from Mayer and Zignago (2011) for any missing observations. As a measure of distance, I use population-weighted distance. The contiguity dummy is equal to one if two countries share a common border. If they do not, it is zero. The language dummy takes the value one if at least 9% of the population speaks the same language and zero otherwise. The colonial dummy indicates whether they were ever in a colonial relationship.

[25] The data to calculate gross output to value added ratios for the manufacturing sector is taken from Inklaar et al. (2011). As this data is biennial, I use the same data point for two consecutive years. Federico's (2004) table D.1. provides data on indices for gross output and value added for the worldwide agricultural sector.

[26] I use the yearly pound sterling per dinar exchange rate in Fouquin and Hugot (2016b) to convert dinars into pound sterling.

A.2 Robustness Appendix

This subsection presents a battery of robustness and sensitivity checks. Many more specifications were tried and I present only the most relevant ones. Table (A.3) shows the estimation results for additional econometric specifications of equation (3.6). Columns (1) - (3) are a reproduction of column (2) in table (1), but include asymmetric pair fixed effects, a time trend interacted with the pair fixed effects, and both. The inclusion of pair-trends causes collinearity with country-pair specific variables. In my specifications most variation over time of the individual country-pairs is already captured by the distance and border variables. In column (2) Stata drops the coefficient on $\ln Dist_{ij,1937}$ and fails to create standard errors for the coefficient on $INTL\_BRDR_{ij,1937}$. In column (3) Stata drops both regressors entirely. Accordingly, the remaining coefficients should be interpreted differently. One way to think about the remaining coefficients on $\ln Dist_{ij,t}$ and $INTL\_BRDR_{ij,t}$ in columns (2) and (3) is that they should be interpreted as deviations from a trend. Given this interpretation, the main results concerning the border and distance elasticities is considered robust.

Columns (4) to (6) show the main specification and the solution to the interwar distance puzzle without using pair fixed effects. Instead, I include the standard gravity covariates *contiguity*, *colonial* and *commonlanguage*. The coefficient on $\ln Dist_{ij,1925}$ in column (4) is not significantly different from the coefficient on $\ln Dist_{ij,1937}$. However, once we exclude sterling bloc and IPS membership from the regression, the distance puzzle reappears.

We also observe a very large and increasing border effect in the regressions without pair fixed effects. The estimate in column (4) for example implies that, all else being equal, international borders decreased trade by an average of $100 * (e^{\beta_{2,1928}} - 1) = 100 * (e^{-4.869} - 1) = 99\%$ in 1928. While this coefficient is larger than comparable coefficients for the present period, it is close to the coefficient estimated by Fouquin and Hugot (2016a) for the interwar period.[27] Given the much higher level of protectionism in the 1920s and 1930s, it is not surprising to find a larger border effect. Yet there might be bilateral factors not controlled for in columns (4) - (6) that bias the absolute value of the coefficients on distance and border. However, for my counterfactual analysis the absolute value is of little relevance if one assumes a constant elasticity of trade to trade costs. What matters is how these individual trade costs change over time and the change in the distance and border variables is not qualitatively different if I include pair fixed effects. For example, multiplying the coefficient on $INTL\_BRDR_{ij,1925}$ in column (4), table (A.3) with the estimate on $INTL\_BRDR_{ij,1937}$ in column (2), table (1) yields a coefficient of -5.285 ($\ln(e^{-4.652} * e^{-0.633})$), which is larger than the -5.151 obtained in column (4) here.

Finally, note that the indicator variables for gold standard adherence, sterling bloc and IPS membership are significantly larger than in the regressions with pair fixed effects. Not controlling for potential endogeneity and unobserved country-pair heterogeneity drastically increases these coefficients, confirming the view of Bergstrand et al. (2015) that not including pair fixed effects biases the coefficient on economic integration agreements.

Table (A.4) presents the results for additional specifications of estimating equation (3.6). Column

[27] Yotov et al. (2016) find a coefficient of -2.474 for the year 2006. Fouquin and Hugot (2016a) do not present results in table form, but their figure 12 suggests that the coefficient on international borders varied between -4.5 and -6 during the interwar period.

Table A.3: Regression Estimates: Robustness

Horse Race

VARIABLES	(1) PPML X_{ij}	(2) PPML X_{ij}	(3) PPML X_{ij}	(4) PPML X_{ij}	(5) PPML X_{ij}	(6) PPML X_{ij}
$\ln(Dist_{ij,1925})$				-0.425***	-0.426***	-0.404***
				(0.0693)	(0.0696)	(0.0686)
$\ln(Dist_{ij,1928})$	-0.0155	-0.0259**	-0.0211*	-0.435***	-0.436***	-0.441***
	(0.0124)	(0.0118)	(0.0121)	(0.0651)	(0.0654)	(0.0665)
$\ln(Dist_{ij,1931})$	-0.0714**	-0.0980***	-0.0898**	-0.535***	-0.496***	-0.502***
	(0.0283)	(0.0375)	(0.0365)	(0.0661)	(0.0701)	(0.0690)
$\ln(Dist_{ij,1934})$	0.0317	-0.0190	-0.0154	-0.438***	-0.335***	-0.348***
	(0.0283)	(0.0153)	(0.0153)	(0.0632)	(0.0729)	(0.0716)
$\ln(Dist_{ij,1937})$	0.0524			-0.426***	-0.318***	-0.309***
	(0.0365)			(0.0676)	(0.0778)	(0.0773)
$INTL\_BRDR_{ij,1925}$				-4.652***	-4.689***	-4.628***
				(0.158)	(0.160)	(0.157)
$INTL\_BRDR_{ij,1928}$	-0.0623*	0.0793**	0.168***	-4.869***	-4.894***	-4.614***
	(0.0365)	(0.0345)	(0.0348)	(0.164)	(0.167)	(0.151)
$INTL\_BRDR_{ij,1931}$	-0.291***	-0.0597	0.116	-5.028***	-5.094***	-4.809***
	(0.0687)	(0.0973)	(0.0949)	(0.152)	(0.159)	(0.154)
$INTL\_BRDR_{ij,1934}$	-0.672***	-0.432***	-0.161***	-5.247***	-5.360***	-5.319***
	(0.0691)	(0.0367)	(0.0371)	(0.150)	(0.165)	(0.161)
$INTL\_BRDR_{ij,1937}$	-0.630***	-0.363		-5.155***	-5.275***	-5.302***
	(0.0925)	(0)		(0.164)	(0.183)	(0.183)
$Contiguity_{ij}$				0.515***	0.528***	0.548***
				(0.0767)	(0.0809)	(0.0833)
$Colonial_{ij}$				0.933***	1.012***	1.001***
				(0.0904)	(0.0945)	(0.0945)
$CommonLanguage_{ij}$				-0.0180	0.0501	0.0746
				(0.0607)	(0.0627)	(0.0616)
$Gold_{ij}$	0.0838**	-0.0142	-0.0132	0.353***	0.334***	
	(0.0333)	(0.0238)	(0.0238)	(0.0763)	(0.0785)	
$SterlingBloc_{ij}$	0.127	0.106	0.123*	0.713***		
	(0.0784)	(0.0714)	(0.0719)	(0.112)		
IPS_{ij}	0.202**	0.418***	0.437***	0.643***		
	(0.0894)	(0.0955)	(0.0935)	(0.153)		
Observations	5,075	5,085	5,075	5,085	5,085	5,085
Country Pair FE's	Yes	Yes	Yes	No	No	No
Asymmetric Pair FE's	Yes	No	Yes	No	No	No
Time trend	No	Yes	Yes	No	No	No

Notes: All estimates are obtained with data for the years 1925, 1928, 1931, 1934 and 1937, and use exporter-time and importer-time fixed effects. The estimates of fixed effects are omitted for brevity. Standard errors are clustered by country pair in parentheses; *** p<0.01, ** p<0.05, * p<0.1

(1) simply reproduces column (7) in table (1) excluding $Dist_{ij,t}$ from the regression. The sterling bloc is now significant at the 10% level and all other coefficients are similar in magnitude to those in column (7) of table (1). Dropping the gold standard dummy (column (2)) reduces the significance of the sterling bloc, but adding three-year lag variables and the gold standard dummy increases its size and significance (column (3)). The coefficient on sterling bloc also remains significant when I use four-year lag variables (column (4)) or add $Dist_{ij,t}$ again to the regression (column (5)). Column (5) in particular is a reproduction of column (2) in table (1) extended to account for phasing-in effects of the bloc dummies. The fact that the sterling bloc is now significant is taken as justification of its use in the trade cost function.

Finally, table (A.5) reproduces the most important regressions with data for all 13 years. Column (1), for example, is the same specification as in columns (1) and (2) in table (1). All coefficients are comparable in size and significance and the significance of the sterling bloc is even increased by adding all years. Moreover, columns (5) to (6) again show that the interwar distance puzzle is resolved once we include IPS and sterling bloc membership. This result is robust to the inclusion of asymmetric pair fixed effects. Again, the inclusion of a time trend causes collinearity problems and the coefficients on $INTL\_BRDR_{ij,1937}$ and $Dist_{ij,1937}$ are dropped in the specifications in columns (3) and (4). The interpretation of the remaining coefficients should change accordingly.

Table A.4: Regression Estimates: Blocs lagged

VARIABLES	(1) PPML X_{ij}	(2) PPML X_{ij}	(3) PPML X_{ij}	(4) PPML X_{ij}	(5) PPML X_{ij}
$\ln(Dist_{ij,1928})$					-0.0181
					(0.0125)
$\ln(Dist_{ij,1931})$					-0.0747**
					(0.0292)
$\ln(Dist_{ij,1934})$					0.0351
					(0.0281)
$\ln(Dist_{ij,1937})$					0.0611
					(0.0378)
$INTL\_BRDR_{ij,1928}$	-0.106***	-0.0533***	-0.127***	-0.103***	-0.0828*
	(0.0201)	(0.0155)	(0.0302)	(0.0223)	(0.0440)
$INTL\_BRDR_{ij,1931}$	-0.447***	-0.396***	-0.491***	-0.435***	-0.338***
	(0.0309)	(0.0276)	(0.0475)	(0.0455)	(0.0803)
$INTL\_BRDR_{ij,1934}$	-0.607***	-0.638***	-0.645***	-0.597***	-0.721***
	(0.0352)	(0.0353)	(0.0391)	(0.0462)	(0.0755)
$INTL\_BRDR_{ij,1937}$	-0.520***	-0.558***	-0.518***	-0.520***	-0.638***
	(0.0392)	(0.0405)	(0.0399)	(0.0398)	(0.0949)
$Gold_{ij}$	0.107***		0.117***	0.104***	0.0996**
	(0.0326)		(0.0378)	(0.0368)	(0.0393)
$Gold_{ij,t-3}$			0.0453		0.0496
			(0.0321)		(0.0315)
$SterlingBloc_{ij,t}$	0.125*	0.0983	0.132*	0.105*	0.163**
	(0.0725)	(0.0675)	(0.0788)	(0.0632)	(0.0763)
$SterlingBloc_{ij,t-3}$		0.0182	-0.00232		-0.0698
		(0.0725)	(0.0739)		(0.0744)
$IPS_{ij,t}$	0.272***	0.272***	0.298***	0.290***	0.267***
	(0.0757)	(0.0884)	(0.0877)	(0.0542)	(0.0929)
$IPSij,t-3$		-0.0124	-0.0307		-0.0653
		(0.0407)	(0.0388)		(0.0483)
$Gold_{ij,t-4}$				-0.0127	
				(0.0408)	
$SterlingBloc_{ij,t-4}$				0.0616	
				(0.0854)	
$IPS_{ij,t-4}$				-0.0627	
				(0.0927)	
Observations	5,085	5,085	5,085	5,085	5,085
Country Pair Fixed Effects	Yes	Yes	Yes	Yes	Yes

Notes: All estimates are obtained with data for the years 1925, 1928, 1931, 1934 and 1937, and use exporter-time, importer-time and pair fixed effects. The estimates of fixed effects are omitted for brevity. Standard errors are clustered by country pair in parentheses; *** p<0.01, ** p<0.05, * p<0.1

Table A.5: Regression Estimates: No Intervall

VARIABLES	(1) PPML X_{ij}	(2) PPML X_{ij}	(3) PPML X_{ij}	(4) PPML X_{ij}	(5) PPML X_{ij}	(6) PPML X_{ij}	(7) PPML X_{ij}
$\ln(Dist_{ij,1926})$	-0.0239***	-0.0175**	-0.0329***	-0.0265***	-0.0188**	-0.0268***	-0.0287***
	(0.00860)	(0.00860)	(0.00838)	(0.00857)	(0.00848)	(0.00848)	(0.00830)
$\ln(Dist_{ij,1927})$	-0.0203**	-0.0159	-0.0260**	-0.0213*	-0.0150	-0.0219**	-0.0199*
	(0.0103)	(0.0102)	(0.0111)	(0.0112)	(0.0102)	(0.0101)	(0.0102)
$\ln(Dist_{ij,1928})$	-0.0192	-0.0144	-0.0301***	-0.0239**	-0.0139	-0.0208*	-0.0215
	(0.0123)	(0.0126)	(0.0113)	(0.0116)	(0.0126)	(0.0123)	(0.0132)
$\ln(Dist_{ij,1929})$	-0.0303**	-0.0257*	-0.0443***	-0.0370**	-0.0250*	-0.0318**	-0.0321**
	(0.0130)	(0.0132)	(0.0160)	(0.0161)	(0.0130)	(0.0129)	(0.0130)
$\ln(Dist_{ij,1930})$	-0.0837***	-0.0777***	-0.107***	-0.0963***	-0.0783***	-0.0865***	-0.0932***
	(0.0216)	(0.0208)	(0.0290)	(0.0280)	(0.0213)	(0.0213)	(0.0208)
$\ln(Dist_{ij,1931})$	-0.0797***	-0.0731***	-0.0956**	-0.0843**	-0.0619**	-0.0697***	-0.0767***
	(0.0278)	(0.0272)	(0.0373)	(0.0361)	(0.0256)	(0.0257)	(0.0249)
$\ln(Dist_{ij,1932})$	-0.0359	-0.0313	-0.0624**	-0.0532**	-0.0242	-0.00268	-0.00247
	(0.0231)	(0.0232)	(0.0246)	(0.0244)	(0.0225)	(0.0217)	(0.0218)
$\ln(Dist_{ij,1933})$	0.00492	0.00649	-0.0319	-0.0257	0.0144	0.0363*	0.0355*
	(0.0222)	(0.0227)	(0.0198)	(0.0199)	(0.0221)	(0.0213)	(0.0215)
$\ln(Dist_{ij,1934})$	0.0290	0.0321	-0.0117	-0.00398	0.0431	0.0657**	0.0636**
	(0.0277)	(0.0282)	(0.0152)	(0.0153)	(0.0280)	(0.0273)	(0.0278)
$\ln(Dist_{ij,1935})$	0.0479	0.0501	0.00304	0.00888	0.0625*	0.0859***	0.0842***
	(0.0320)	(0.0328)	(0.0119)	(0.0119)	(0.0326)	(0.0318)	(0.0325)
$\ln(Dist_{ij,1936})$	0.0418	0.0436	-0.00631	-0.00257	0.0564	0.0805**	0.0804*
	(0.0400)	(0.0405)	(0.0106)	(0.0108)	(0.0405)	(0.0405)	(0.0414)
$\ln(Dist_{ij,1937})$	0.0531	0.0541			0.0673*	0.0916**	0.0922**
	(0.0365)	(0.0366)			(0.0369)	(0.0362)	(0.0372)
$INTL\_BRDR_{ij,1926}$	-0.0263	-0.0390*	0.0568***	0.0467**	-0.0365*	-0.0188	-0.00904
	(0.0216)	(0.0216)	(0.0213)	(0.0216)	(0.0213)	(0.0213)	(0.0204)
$INTL\_BRDR_{ij,1927}$	-0.0307	-0.0385	0.107***	0.103***	-0.0406	-0.0222	-0.00791
	(0.0271)	(0.0269)	(0.0289)	(0.0288)	(0.0271)	(0.0269)	(0.0248)
$INTL\_BRDR_{ij,1928}$	-0.0476	-0.0558	0.160***	0.155***	-0.0573*	-0.0366	-0.00863
	(0.0341)	(0.0344)	(0.0317)	(0.0321)	(0.0348)	(0.0350)	(0.0335)
$INTL\_BRDR_{ij,1929}$	-0.0258	-0.0332	0.237***	0.232***	-0.0354	-0.0149	0.0122
	(0.0342)	(0.0342)	(0.0441)	(0.0436)	(0.0345)	(0.0345)	(0.0322)
$INTL\_BRDR_{ij,1930}$	-0.104**	-0.114**	0.221***	0.212***	-0.114**	-0.0917*	-0.0544
	(0.0517)	(0.0501)	(0.0757)	(0.0734)	(0.0511)	(0.0513)	(0.0477)
$INTL\_BRDR_{ij,1931}$	-0.267***	-0.279***	0.0997	0.0891	-0.292***	-0.269***	-0.230***
	(0.0645)	(0.0627)	(0.0975)	(0.0949)	(0.0615)	(0.0617)	(0.0587)
$INTL\_BRDR_{ij,1932}$	-0.509***	-0.521***	-0.126**	-0.134**	-0.529***	-0.547***	-0.560***
	(0.0536)	(0.0528)	(0.0611)	(0.0602)	(0.0528)	(0.0534)	(0.0518)
$INTL\_BRDR_{ij,1933}$	-0.604***	-0.610***	-0.148***	-0.151***	-0.614***	-0.632***	-0.645***
	(0.0533)	(0.0533)	(0.0506)	(0.0503)	(0.0537)	(0.0546)	(0.0542)
$INTL\_BRDR_{ij,1934}$	-0.671***	-0.680***	-0.162***	-0.169***	-0.687***	-0.707***	-0.720***
	(0.0683)	(0.0689)	(0.0374)	(0.0375)	(0.0698)	(0.0710)	(0.0719)
$INTL\_BRDR_{ij,1935}$	-0.759***	-0.766***	-0.201***	-0.206***	-0.775***	-0.797***	-0.811***
	(0.0798)	(0.0812)	(0.0273)	(0.0275)	(0.0816)	(0.0829)	(0.0848)
$INTL\_BRDR_{ij,1936}$	-0.742***	-0.749***	-0.141***	-0.145***	-0.759***	-0.783***	-0.802***
	(0.103)	(0.104)	(0.0256)	(0.0263)	(0.105)	(0.107)	(0.109)
$INTL\_BRDR_{ij,1937}$	-0.638***	-0.643***			-0.654***	-0.679***	-0.701***
	(0.0928)	(0.0928)			(0.0942)	(0.0961)	(0.0988)
$Gold_{ij,t}$	0.0691**	0.0657**	0.0121	0.0127	0.0671**	0.0545*	
	(0.0272)	(0.0265)	(0.0172)	(0.0171)	(0.0273)	(0.0279)	
$SterlingBloc_{ij,t}$	0.129*	0.132*	0.0188	0.0204			
	(0.0722)	(0.0764)	(0.0506)	(0.0522)			
$IPS_{ij,t}$	0.202**	0.209**	0.327***	0.332***	0.271***		
	(0.0964)	(0.0943)	(0.0716)	(0.0718)	(0.0819)		
Observations	13,143	13,130	13,143	13,130	13,143	13,143	13,143
Asymetric Pair FE's	No	Yes	No	Yes	No	No	No
Time trend	No	No	Yes	Yes	No	No	No

Notes: All estimates are obtained with data for all years from 1925 to 1937, and use exporter-time, importer-time and pair fixed effects. The estimates of fixed effects are omitted for brevity. Robust standard errors, clustered by country pair, are in parentheses; *** p<0.01, ** p<0.05, * p<0.1

A.3 General Equilibrium Appendix

This section briefly explains the techniques used to calculate GE effects. A more detailed discussion is found in Yotov et al. (2016) or in the Stata code that accompanies this paper (available upon request).

After having implemented the two-stage procedure by Anderson and Yotov (2016) and having obtained the complete set of bilateral trade costs $\tau_{ij,1937}^{BLN}$, I estimate the gravity model in equation (3.1) as follows:

$$X_{ij,1937} = exp[\ln \tau_{ij,1937}^{BLN} + \gamma_i^{BLN} + \delta_j^{BLN}] \tag{A.2}$$

The estimates of the importer fixed effects (δ_j^{BLN}) and of the exporter fixed effects (γ_i^{BLN}) from the above estimation are used to construct the baseline multilateral resistances (MR):

$$[\hat{P}_j^{1-\sigma}]^{BLN} = \frac{E_j}{exp(\hat{\delta}_j^{BLN})} * \frac{1}{E_{DEU}} \tag{A.3}$$

$$[\hat{\Pi}_i^{1-\sigma}]^{BLN} = \frac{Y_i}{exp(\hat{\gamma}_i^{BLN})} * E_{DEU} \tag{A.4}$$

where, output is constructed as $Y_i = \sum_{j=1}^{N} X_{ij}$ and expenditure is calculated as $E_j = \sum_{i=1}^{N} X_{ij}$. E_{DEU} is expenditure of the reference country, Germany, for which the inward MR is normalized to one and the corresponding fixed effect δ_{DEU} is removed from regression (A.2). The predicted volume of trade from regression (A.2) is used to calculate country i's baseline trade volume ($\hat{X}_i = \sum_{j=1}^{N} \hat{X}_{ij}$ for all $j \neq i$). The computation for the counterfactual MRs and trade volumes is analogous.[28] The conditional GE effects are calculated as the difference, in percentage, between the baseline and the counterfactual trade volumes.

The full endowment GE effects are obtained by implementing a four-step iterative procedure. First, I use the market-clearing condition in (3.4) to translate the conditional GE effects on the multilateral resistance terms into first-order changes in factory-gate prices, by applying the definition of the estimated exporter fixed effects in equation (A.4):

$$\Delta p_i^{CF} = \frac{p_i^{CF}}{p_i} = \left(\frac{exp(\hat{\gamma}_i^{CF})/E_{DEU}^{CF}}{exp(\hat{\gamma}_i^{BLN})/E_{DEU}}\right)^{\frac{1}{1-\sigma}} \tag{A.5}$$

In the second step output and expenditure respond endogenously to the above change in factory gate prices: $Y_i^{CF} = (p_i^{CF}/p_i)Y_i^{BLN}$ and $E_j^{CF} = (p_j^{CF}/p_j)E_j^{BLN}$. This in turn will trigger additional changes in the multilateral resistance terms and so forth. The structural gravity equation (3.1) translates the changes in output and expenditure into changes in trade flows:

$$X_{ij,1937}^{CF} = \frac{(\tau_{ij,1937}^{CF})^{1-\sigma}}{(\tau_{ij,1937}^{BLN})^{1-\sigma}} * \frac{Y_i^{CF} E_j^{CF}}{Y_i^{BLN} E_j^{BLN}} * \frac{[\hat{\Pi}_i^{1-\sigma}]^{BLN}}{[\hat{\Pi}_i^{1-\sigma}]^{CF}} * \frac{[\hat{P}_j^{1-\sigma}]^{BLN}}{[\hat{P}_j^{1-\sigma}]^{CF}} * \hat{X}_{ij} \tag{A.6}$$

Equation (A.6) computes a counterfactual value of trade that accounts for changes in output and expenditure, via a change in the factory gate price, and changes in inward and outward multilateral

[28] Note that for the calculation of the counterfactual MRs the original data on output and expenditure is used. The conditional GE values of the MRs under the counterfactual scenario are then calculated analogous to equations (A.3) and (A.4) but use the fixed effects $\hat{\gamma}_i^{CF}$ and $\hat{\delta}_j^{CF}$ from estimating equation (3.9).

resistances. Yet, these changes are only first-order changes, because they only capture the changes in the conditional outward multilateral resistances and the immediate response in the factory-gate prices.

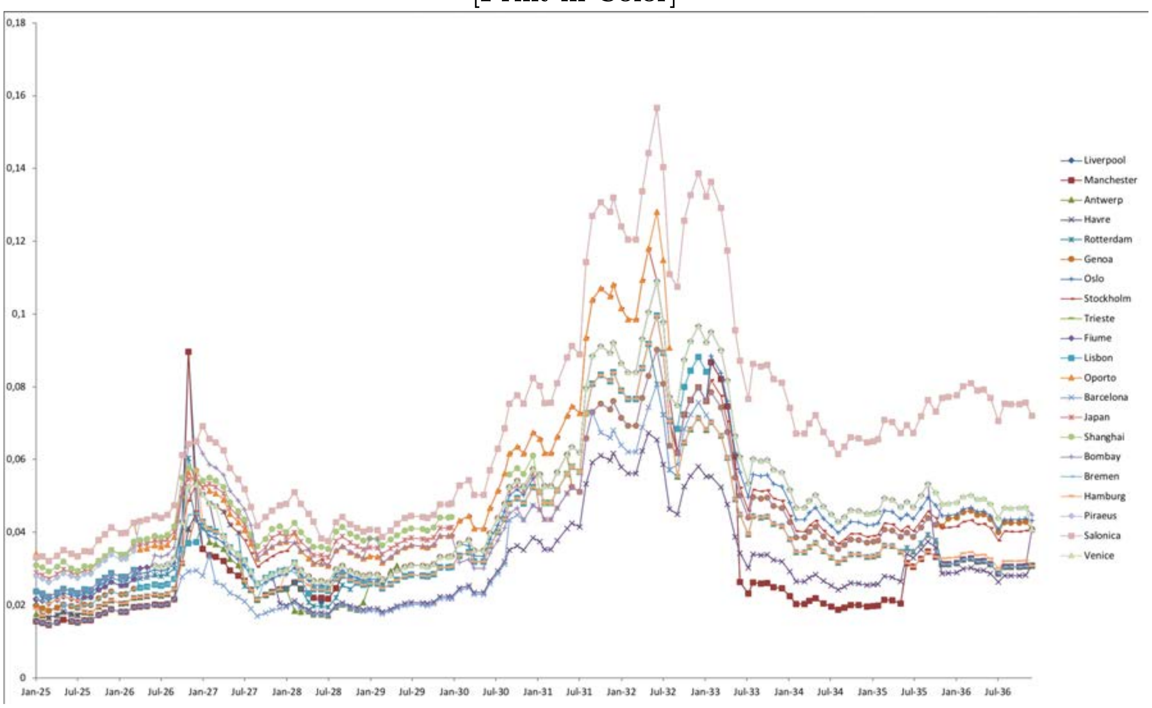
Hence, the third stage of the loop reestimates the gravity model (3.9) with the new value of bilateral trade, $X_{ij,1937}^{CF}$ from equation (A.6), and then computes the corresponding GE effects associated with the new fixed effect estimates. The idea is to update the value of bilateral trade to obtain additional responses in the multilateral resistances and in the values of output and expenditure. Once the new set of fixed effects associated with the new value of trade from equation (A.6) are estimated, the loop starts again at the first stage of the iterative procedure in order to obtain a new set of factory gate prices associated with these fixed effects. These three steps are repeated until the change in each of the factory gate prices is close to zero and the model has reached its new equilibrium. The difference in percentage between the baseline and the new equilibrium trade volumes yields the full endowment GE effect.

A.4 Transport Costs Appendix

Figure (A.2) plots the real freight rates of all 21 routes over time. In general, all routes move together, which is due to nominal freight rates changing very slowly and in tandem across routes. Most of the changes in real freight rates is then caused by the change in the price of cotton. This indicates that there has indeed been a significant degree of cartelization in the liner industry. As discussed above, tramp and liner shipping differ in their way of operation. Prices for tramp shipping are usually set in spot markets and tramps are hired on a charter basis. Liners run on fixed routes and fixed timetables, which makes the industry potentially more susceptible to cartelization. Indeed, the liner industry of the post-war period is organized into conferences, which discuss, and perhaps collude in, setting prices and market shares (Hummels, 2007).

Figure A.2: Cotton Ad-Valorem Freight Rates from New York to 21 Destinations

[Print in Color]



Notes: The graph shows cotton (American middling) freight rates for 21 routes deflated by the price at the place of origin (New York). *Sources*: Commercial and Financial Chronicle (1925 - 1936) and Statistisches Reichsamt (1936).

Following Hummels (1999, 2007) and Estevadeordal et al. (2003) I can estimate the technological relationship between distance and transport costs with a simple OLS regression

$$\ln\left(f_{n,t}\right) = \alpha + \beta \ln\left(Dist_{n,t}\right) + D_t + \epsilon_{n,t} \tag{A.7}$$

where $f_{n,t}$ is the ad-valorem freight, $Dist_{n,t}$ is great circle distance between ports and D_t is an optional time dummy. Table (A.6) shows the regression results with and without time dummies for all years and for the pre and post-depression periods.

The coefficients on distance in column (1) and (2) are significantly larger than those obtained

Table A.6: Technology in the Interwar Liner Shipping Industry

	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	OLS	OLS	OLS	OLS	OLS
VARIABLES	$\ln(f_{n,t})$	$\ln(f_{n,t})$	$\ln(f_{n,t})$	$\ln(f_{n,t})$	$\ln(f_{n,t})$	$\ln(f_{n,t})$
$\ln(Dist_{n,t})$	0.397***	0.499***	0.415***	0.467***	0.320***	0.440***
	(0.0374)	(0.0186)	(0.0482)	(0.0280)	(0.0436)	(0.0188)
Constant	-6.530***	-7.390***	-6.466***	-7.123***	-5.887***	-6.742***
	(0.312)	(0.164)	(0.401)	(0.240)	(0.363)	(0.164)
Observations	2,751	2,751	1,635	1,635	2,204	2,204
R-squared	0.039	0.778	0.043	0.697	0.024	0.829
Time Dummy	No	Yes	No	Yes	No	Yes
Years	All	All	1927-1929	1927-1929	1934-1936	1934-1936

Notes: The distance between New York and 21 port cities is taken from https://www.distance-cities.com/
Standard errors in parentheses; *** p<0.01, ** p<0.05, * p<0.1

by Hummels (1999, 2007) for the post-war period, which one would expect given the technological improvements such as containerization.[29] When I estimate equation (A.7) for the period between May 1927 and May 1929 the elasticity of ocean transportation to distance is 0.415 (column (3)), but drops to 0.32 for the period May 1934 and May 1936. This drop is still apparent, although somewhat smaller when I include time fixed effects. This result suggests that, if anything, there was technological improvement in the shipping sector during the interwar period.

[29] Note, however, that Hummels (1999, 2007) includes a value to weight ratio in his regressions. This is not possible here, since I only consider one good.

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