

Has Distance Died? Evidence from a Panel Gravity Model

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The estimated coefficient of distance on the volume of trade is generally found to increase rather than decrease through time using the traditional gravity model of trade. This distance puzzle proved robust to several ad hoc versions of the model using data for 1962–96 for a large sample of 130 countries. The introduction of an “augmented” barrier to trade function removes the paradox, yielding a decline in the estimate of the elasticity of trade to distance of about 11 percent over the 35-year period for the whole sample. However, the “death of distance” is shown to be largely confined to bilateral trade between rich countries, with poor countries becoming marginalized.

There is a widespread perception that the current wave of globalization, much like the first, should have led to the “death of distance.” Other things equal, globalization should generate a dispersion of economic activity reflecting a decline in transaction costs, especially transport costs. But studies based on the traditional gravity model of international trade—the workhorse for studies on the pattern of trade and the influence of transport costs—do not reach that conclusion. Although some studies based on the gravity model have used direct

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THE WORLD BANK ECONOMIC REVIEW

doi:10.1093/wber/lhi004

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measures of transport cost barriers to trade,¹ most rely on distance as a proxy for transport costs, obtaining an estimated elasticity of bilateral trade with respect to distance in the range $[-1.3; -0.8]$. However, as will be detailed, when the model is estimated separately for several years, the absolute value of the coefficient almost always increases over time. This is puzzling, because the common perception of globalization is that distance should be becoming less important in international trade, implying decreasing rather than increasing values for the estimated coefficient of distance.

This paradoxical result was initially investigated by Brun and others (1999) in a traditional gravity model framework. Earlier, Leamer and Levinsohn (1995, pp. 1387–88), reviewing the literature on international trade and distance, noted that “the effect of distance on trade patterns is not diminishing over time. Contrary to popular impression, the world is not getting dramatically smaller.” They conclude that “dispersion of economic mass is the answer, not a shrinking globe” for this result. In a recent examination of the paradox Coe and others (2002) review explanations in the literature. One is the exclusion of zero observations from the model, which could bias estimation of the impact of distance over time because of the changing composition of trade. Another is that the traditional gravity model omits what is now being referred to as “multilateral trade resistance.” Both explanations are tested here. The paradoxical result remains. A third, obvious explanation is examined here as well: the possibility of misspecification of the transport cost function due to omitted-variable bias.

Each of these potential explanations is examined for a large sample of 130 countries spanning bilateral trade over 35 years. The contribution of this study comes from the combination of more appropriate estimation techniques and a more thorough treatment of transport costs. The article first considers whether the puzzle still holds when estimation is carried out in a random-effects panel procedure with correction for endogeneity (Hausman and Taylor estimator) and for potential selection bias. The puzzle is found to be a robust result in a large sample. Next, a better specified gravity equation is considered. To bring it closer to accepted theoretical foundations, it includes a measure of remoteness in the estimates. Most important, it also defines an “augmented” transport cost function that includes indexes of infrastructure, price of oil, and composition of trade as arguments, all of which turn out to be statistically significant along

1. Some recent studies use direct measures of transport costs in a gravity model. Baier and Bergstrand (2001) use c.i.f.-f.o.b. price ratios over 1970–85 for 16 Organisation for Economic Co-operation and Development countries. Arguing that such comparisons provide limited information (Hummels and Lugovskyy 2003), Hummels (1999, 2001a, b) has studied determinants of freight costs based on more reliable customs data (such as German shipping data over the past 40 years or U.S. customs data at the five-digit SITC level over the period 1974–98). In the same way, Limao and Venables (2001) estimate maritime and land transport costs using 1990 transport cost data for container shipments from Baltimore to 64 destination cities, and Estevadeordal and others (2003) use British maritime transport cost data for a sample of 46 countries over the period 1870–1938.

predicted lines and contribute to the solution of the puzzle. To check for robustness, the sample is split into low-income countries and high-income countries.

To anticipate the main results: Once remoteness is included in the model, the increasing impact of distance on trade vanishes, and with the introduction of an augmented transport cost function into the log-linear specification of the gravity model, the absolute value of the elasticity of bilateral trade with respect to distance decreases significantly over time. Furthermore, when the sample is split into low- and high-income countries, the elasticity of bilateral trade with respect to distance reveals no trend for low-income countries' trade, whereas it falls for bilateral trade between high-income countries, a result in accordance with estimates obtained using more reliable customs data on transport cost (available for only a handful of countries). This result may reflect the fact that low-income countries have been marginalized in the current wave of globalization.

I. IS THERE A PUZZLE?

The puzzle is found in the results for the traditional gravity model using cross-sectional data. This section examines whether the puzzle remains when panel data for a sample of 130 countries for 1962–96 are used. It also addresses issues raised by alternative treatments of missing observations and by potential selection bias. The bulk of the literature that reaches the conclusion that the coefficient for distance increases in absolute value over time is based on the traditional gravity model. Estimated in cross-section, this model is a variant of the following log-linear equation:

$$(1) \quad \ln(T_{ij}) = \alpha_0 + \alpha_1 \ln(Y_i Y_j) + \alpha_2 \ln(N_i N_j) + \beta \ln(D_{ij}) + \alpha_3 DUM_{ij}^k + \varepsilon_{ij},$$

where T_{ij} is the total volume of trade between country i and country j , which depends on the product of partners' income, $Y_i Y_j$, the product of partners' respective populations, $N_i N_j$, the distance between i and j , D_{ij} , and a vector of dummy variables, DUM_{ij}^k . These dummy variables usually capture a common language, a common land border, a common colonizer, the condition of being landlocked, the existence of a free trade area, and sometimes a common currency.

Typically, equation 1 is estimated in a cross-section setting for different years (with or without importer and exporter fixed effects). The puzzle is revealed in a negative impact of distance ($\beta < 0$) that increases absolutely through time. For example, in a sample of 74 countries for 1965 and 1992, Frankel (1997, table 4.2) obtains $\beta^{65} = -0.48$ and $\beta^{92} = -0.77$. Likewise, Leamer (1993) estimates distance elasticities, which did not fall between 1970 and 1985. As pointed out by Coe and others (2002), several other studies have failed to find a declining coefficient for distance over time, and most have found a significant increase in the absolute value of the estimated coefficient. For

instance, estimating equation 1 over a sample of 130 countries for 1962 and 1996 yields: $\beta^{62} = -0.86$ and $\beta^{96} = -1.34$, representing an increase of the impact of distance on bilateral trade of about 36 percent over 35 years.²

Setting the theoretical underpinnings of the gravity equation aside until the following section, this puzzling result is put to the scrutiny of typical econometric problems. Estimated in cross-section, equation 1 has several shortcomings. First, because the dummy variables capture only part of the unobservable heterogeneity of country pairs, the remaining unobservable heterogeneity could potentially bias estimates of the coefficient for distance. Second, the typical ordinary least squares estimates may be prone to omitted-variables bias. Third, there is a potential selection bias due to missing values in bilateral trade data.

The use of panel data, with a time dimension in addition to the traditional importer and exporter dimensions, can address the issue of unobservable heterogeneity of country pairs. The usual correction introduces three specific effects: exporter, importer, and time effects (see Matyas 1997; Soloaga and Winters 2001). But the model with three specific effects is only a restricted version of the more general model adopted here, which allows for country-pair heterogeneity (see Cheng and Wall 1999; Egger and Pfaffermayr 2003). These bilateral specific effects are included to capture all unobservable time-invariant characteristics of the bilateral trade relationships that might otherwise be captured by the distance coefficient.

Moreover, because the focus is on the death of distance, panel data with a time dimension allow estimating a time-varying elasticity of trade with respect to distance. Thus, β is allowed to change over time but not across countries (differences across countries are permitted in section III, which splits the sample into groups). Finally, a quadratic term is used to allow for the existence of a turning point:

$$(2) \quad \ln(M_{ijt}) = \alpha_0 + \alpha_1 \ln Y_{it} + \alpha_2 \ln Y_{jt} + \alpha_3 \ln N_{it} + \alpha_4 \ln N_{jt} + \beta_0 \cdot t + \beta_1 \cdot \ln D_{ij} \\ + \beta_2 \cdot t \cdot \ln D_{ij} + \beta_3 \cdot t^2 \cdot \ln D_{ij} + \varepsilon_{ijt} = Z_1 \psi_1 + \beta_t \cdot \ln D_{ij} + \varepsilon_{ijt},$$

where t is a time trend, $\varepsilon_{ijt} = \mu_{ij} + v_{ijt}$ with μ_{ij} is a specific bilateral random effect, and v_{ijt} is the idiosyncratic error term with the usual properties.

In this equation, used as a starting point for examining the puzzle, bilateral imports (of i from j) rather than total bilateral trade (M_{ij} instead of T_{ij}) are used as the dependent variable, and a random-effects estimation procedure is used to avoid eliminating the coefficient for distance in the equation (the within-transformation in a bilateral fixed-effects model removes variables that are cross-sectional

2. Appendix C reports the evolution of the distance coefficient (equation C1) when equation 1 is estimated for each year with country fixed effects as in the recent literature (Rose and Van Wincoop 2001; Anderson and van Wincoop 2003; Feenstra 2003).

time-invariant). The cost of moving to a random-effects estimation procedure is that some explanatory variables should be correlated with the bilateral random effects, a potential problem addressed later. An instrumental procedure is used to correct for this potential endogeneity (see appendix B).

The data set covers 130 countries for the years 1962–96, a period that captures most of the current wave of globalization during which transport costs have purportedly fallen. Import trade statistics are taken from the United Nations Commodity Trade Statistics (Comtrade) database.

Although the usual restriction of equal coefficients for origin and destination countries is rejected by the data (especially for the population coefficients), the difference in specification has no effect on the values of the parameters of interest β_1 , β_2 , and β_3 (table 1, columns 1 and 2).³ The impact of distance on trade increases over time, because $|\beta_t| \sim 1.321 + (0.0052.t) - (0.0001.t^2)$ (table 1, column 2). Thus estimation on panel data including bilateral specific effects does not resolve the puzzle observed in the literature based on cross-section data.

However, endogeneity tests find the gross domestic product (GDP) variables to be endogenous—correlated with the bilateral specific effects (see details in appendix B). It is therefore sensible to take the results corresponding to equation 2 estimated with the instrumental variables estimator proposed by Hausman and Taylor (1981) as typical of the results for a gravity trade model using distance as a proxy for transport costs (table 1, column 3). The overall fit is good (R -squared = 0.61), and all variables have the expected sign and plausible values. As suggested by theory, the elasticity of trade to income is significant and close to unity. The population variables have the expected negative sign, capturing the often observed phenomenon that trade tends to constitute a smaller percentage of GDP for larger countries, as discussed later. The conclusion from this preliminary inquiry is that the puzzle persists, yielding the following estimate of the evolution of the absolute value of the coefficient for distance:

$$|\beta_t| \sim 1.268 + (0.0062.t) - (0.0001.t^2).$$

This evolution of $|\beta_t|$ over 1962–96, as estimated in equation 2, is plotted in figure 1. These estimates indicate that a 10 percent increase in distance would reduce bilateral trade by 12.7 percent in 1962 and by 13.8 percent in 1996, or an 8.7 percent increase in the impact of distance over 35 years instead of the expected decrease.

How robust is this result? The first concern is the large number of missing values. The sample has a potential of $130 \times 129 \times 35 = 586,950$ observations.

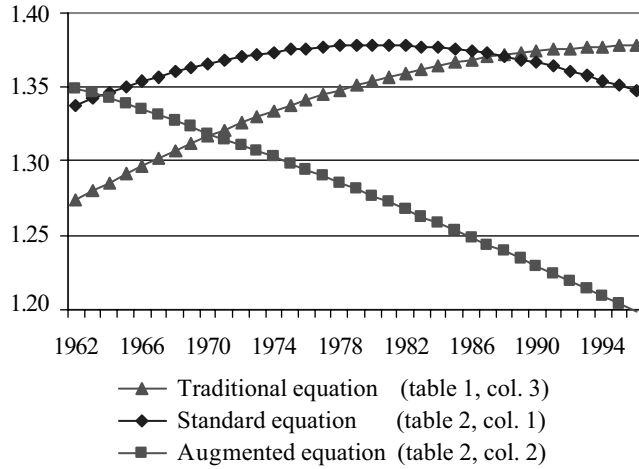
3. Several recent studies use panel data without a time-series dimension but with country fixed effects (for origin countries and destination partners). This estimation procedure forces them to impose identical coefficient values on the income and population variables (Rose and Van Wincoop 2001; Coe and others 2002; Anderson and van Wincoop 2003).

TABLE 1. Distance in a Traditional Panel Gravity Model

Variable	1		2		3		4		5	
	GLS		GLS		HT		HT (with zero values)		HT (with variables for selection bias)	
$\ln Y_{it}$			1.000 (98.76)		0.833 (39.80)		0.764 (43.13)		0.808 (40.77)	
$\ln Y_{jt}$			1.159 (115.87)		1.218 (54.81)		1.571 (104.42)		1.057 (46.82)	
$\ln(Y_{it}Y_{jt})$	1.080 (141.55)									
$\ln N_{it}$			-0.076 (5.95)		-0.049 (2.44)		-0.057 (4.90)		-0.022 (1.15)	
$\ln N_{jt}$			-0.164 (12.90)		-0.251 (12.51)		-0.554 (40.85)		-0.175 (8.81)	
$\ln(N_{it}N_{jt})$	-0.121 (13.30)									
$\ln D_{ij}$	-1.320 (51.90)		-1.321 (52.17)		-1.268 (70.93)		-1.309 (76.40)		-1.203 (69.48)	
t	-0.027 (6.11)		-0.027 (6.01)		-0.030 (5.14)		-0.044 (8.32)		-0.011 (1.91)	
$t \cdot \ln D_{ij}$	-0.0052 (9.52)		-0.0052 (9.63)		-0.0062 (9.10)		-0.0043 (6.59)		-0.0064 (9.49)	
$t^2 \cdot \ln D_{ij}$	0.00010 (22.04)		0.00010 (22.06)		0.00009 (14.76)		0.00009 (16.04)		0.00010 (16.95)	
No. observations	171,998		171,998		171,998		216,511		171,998	
R-squared	0.58		0.59		0.61		0.59		0.64	
Hausman test	401.62 $\chi^2(5)$		619.85 $\chi^2(7)$							
GLS vs. HT					9666.27 $\chi^2(8)$		13,727.05 $\chi^2(8)$		2418.34 $\chi^2(8)$	

Source: Authors' computations based on data described in appendix A.

Note: Numbers in parentheses are t -statistics. GLS is generalized least squares estimator and HT is Hausman and Taylor (1981) estimator.

FIGURE 1. The Elasticity of Bilateral Trade to Distance $|\beta_t|$ 

Source: Authors' computations based on data described in appendix A.

Missing values are reported for 71 percent of the potential observations, resulting in 171,998 observations. The database does not distinguish between countries that do not report their trade statistics (missing values) and country pairs with no bilateral trade (zero trade). Some country pairs report missing values at the beginning of the sample period and positive trade at the end. For them, the missing values can be suspected of being zeroes (perhaps because transport costs are too high at the beginning of the sample period). Excluding these zeroes would bias the results. The missing values are assumed to equal zero if they occur before 1975, provided that positive trade is observed thereafter. As in Frankel (1997, p. 145), zeroes are replaced by $M_{ijt}=1$. The total number of observations rises by 26 percent. Results under this specification do not change the increasing impact of distance on trade (see table 1, column 4).

The missing values are a source of potential selection bias. Following Nijman and Verbeek (1992), three additional explanatory variables are introduced into equation 2 to correct for the selection bias (see table 1, column 5).⁴ Even if two of these proxies are significant, the estimates for β_1 , β_2 , and β_3 remain similar.

Estimates of the evolution of β_t could also be biased because some series may contain a unit root, in which case the estimates in table 1 would be spurious if the relations were not cointegrated. So, a Levin and Lin (1993) unit root test

4. The following variables are added in the equation (coefficient; t Student): the number of years of presence of the couple ij in the sample (0.04; 31.98); a dummy variable that takes the value 1 if ij is observed during the entire period (0.016; 0.74); and a dummy variable that takes the value 1 if ij is present in $t-1$ (0.54; 35.02). Together, these variables have been shown in Monte Carlo experiments by Nijman and Verbeek (1992) to be good proxies for the Heckman correction term for selection bias.

was applied to the series for GDP, population, and bilateral imports. This test rejects, very significantly, the null hypothesis of a unit root for all series.

Finally, to check that the increasing impact of distance on trade does not capture tendencies in other coefficients, two sets of regressions were estimated: year-by-year regressions with country fixed effects and regressions over three-year subperiods with bilateral effects and the Hausman and Taylor (1981) estimator. The estimated coefficients from 1962 to 1996, plotted in appendix figure C1, show that the increasing impact of distance over time is unaffected.

From these robustness checks it can be concluded that the estimated value of $|\beta_i|$ increases over time in the traditional gravity model. The puzzle remains.

II. EXPLAINING THE PUZZLE

The puzzle could be the result of a misspecification of the traditional gravity equation. A more solid theoretical foundation for the gravity equation is found in the currently popular application of the gravity model to aggregate trade between economies assumed to be specialized in differentiated products. As shown by Deardoff (1998) and Anderson and Van Wincoop (2003, 2004), utility maximization of an identical constant elasticity of substitution (CES) utility function (over countries) yields the following expression for the cost, insurance, and freight (c.i.f.) value of bilateral imports,

$$(3) \quad M_{ijt} = (Y_{it}Y_{jt}/Y_w)(\theta_{ijt}/\tilde{P}_{it}\tilde{P}_{jt})^{1-\sigma},$$

where Y_w is world income, θ_{ijt} is bilateral transport costs, σ is the elasticity of substitution in the CES utility function, and \tilde{P}_{it} , \tilde{P}_{jt} can be interpreted as multilateral trade resistance indexes whose values are given by:

$$(4) \quad \tilde{P}_{it}^{1-\sigma} = \sum_j (Y_{jt}/Y_w)(\theta_{ijt}/\tilde{P}_{jt})^{1-\sigma}.$$

Estimation of the system of equations 3 and 4 raises issues related to estimation of the multilateral trade resistance index and estimation of transport costs. For the trade resistance index, suppose, as in the literature, that estimation proceeds on a cross-section basis. In that case either data for the price indexes are available (as in Baier and Bergstrand 2001), or \tilde{P}_i and \tilde{P}_j can be estimated directly (from the structural model with nonlinear least squares), following Anderson and Van Wincoop (2003). Alternatively, as explained by Anderson and Van Wincoop (2003) and Feenstra (2003), country-specific effects can be used to capture the variation in \tilde{P}_i and \tilde{P}_j .⁵

Thus, a first method of estimating the gravity equation uses panel data techniques relying on a cross-section specification with a fixed-effects estimator

5. See the discussion by Anderson and Van Wincoop (2004, p. 712) on the three ways to estimate the theoretical gravity model on a cross-section basis.

that includes importer and exporter dummy variables (Rose and Van Wincoop 2001; Coe and others 2002; Anderson and Van Wincoop 2003). However, the preferred approach, used here, relies on a random-effects estimator, with the two dimensions being country-pair and time-specific effects. Of these two panel techniques the country fixed-effects estimator offers more variability on the distance coefficient because it is estimated year by year. But the risk with this method is that the distance coefficient will capture all unobserved bilateral characteristics that cannot be included in the specification (other than the usual dummy variables for common language, common border, and so on). The country-pair random effects estimator with a time dimension controls for all the unobserved bilateral specific effects, but at the cost of imposing a trend specification for the evolution of distance. After weighing the tradeoffs, the country-pair random-effects estimator is selected (but for illustrative purpose, appendix C reports the results for year-by-year estimations with country fixed effects). The choice of this method requires replacing the multilateral trade resistance indexes with a remoteness measure.⁶ This implies substituting values of R_{it} and R_{jt} for \tilde{P}_{it} and \tilde{P}_{jt} in equation 3.

Estimation of bilateral transport costs, θ_{ijt} , starts with the standard transport cost function in which distance reflects marginal transport costs. (Later, an augmented version is proposed.) In the standard implementation, used by Hummels (2001a) and Anderson and Van Wincoop (2003) among others, transport costs include distance (D_{ij}) and a vector of dummy variables for common border (B_{ij}) and being landlocked ($L_{i(j)}$). Assuming the standard multiplicative form yields:

$$(5) \quad \theta_{ijt} = (D_{ij})^{\gamma_t} e^{\delta_1 \cdot B_{ij} + \delta_2 L_i + \delta_3 L_j}$$

with expected signs $\delta_1 < 0$, $\delta_2 > 0$, and $\delta_3 > 0$. Once more, because bilateral specific effects capture the time-invariant characteristics of bilateral trade, equation 5 is estimated without including the dummy variables. The elasticity of transport costs to distance, γ_t , is assumed to be approximated by a quadratic time trend (t):

$$(6) \quad \gamma_t \equiv ([\partial \theta_{ijt} / \theta_{ijt}] / [\partial D_{ij} / D_{ij}]) = \gamma_1 + \gamma_2 \cdot t + \gamma_3 \cdot t^2.$$

6. As explained, the literature on estimating the theoretical gravity model, notably using multilateral trade resistance indexes, is based on cross-section data. With a time dimension added in the estimation, none of the three suggestions is applicable: Price indexes are not available; the structural estimation method proposed by Anderson and Van Wincoop (2003) is not suitable for three dimensions (importer, exporter, and time); and country dummy variables no longer capture the multilateral resistance indexes, which change over time. Hence, the multilateral trade resistance indexes are replaced by a remoteness measure (as in Soloaga and Winters 2001) even if, as Anderson and Van Wincoop (2003) show, the functional form of the remoteness variable is not in conformity with the theory and can introduce a bias. However, it is comforting that the year by year estimations with country fixed effects—without bias due to the remoteness variables—reported in appendix C yield similar results.

Finally, in estimating a gravity model using panel data with a long time dimension (35 years in this case), it is essential to capture the effects of changes in relative prices, because a normalization of prices to unity can no longer be justified. For a large sample of countries for which representative price indexes are not available, real exchange rate indexes have to be used. As in Soloaga and Winters (2001) and Bayoumi and Eichengreen (1997), among others, the bilateral real exchange rate between i and j , RER_{ijt} is introduced into the equation. Also, RER_{ijt} may be interpreted as a proxy for unobservable movement of multilateral resistance indexes through time.⁷

Under these working assumptions, estimation of the standard trade barrier function in the gravity model boils down to plugging equation 5 into a modified version of equation 3 that also includes population to proxy Engel effects and, on the supply side, differences in factor endowments (Bergstrand 1989; Frankel 1997; Soloaga and Winters 2001; Coe and others 2002).⁸ This yields an equation similar to equation 2, except for the inclusion of “remoteness” that takes into account relative transport costs, and the real exchange rate that takes into account the effects of the evolution in relative prices. This results in:

$$(7) \quad \ln(M_{ijt}) = Z_1\psi_1 + \beta_t \ln D_{ij} + \alpha_5 \ln R_{it} + \alpha_6 \ln R_{jt} + \alpha_7 \ln RER_{ijt} + \varepsilon_{ijt} \\ = Z_1\psi_1 + Z_2\psi_2 + \beta_t \ln D_{ij} + \varepsilon_{ijt}$$

with $\alpha_5 > 0$, $\alpha_6 > 0$, $\alpha_7 < 0$, and according to equations 3 and 5, $\beta_t = (1 - \sigma)\gamma_t > 0$. Estimation results for equation 7 are reported in table 2, column 1, and should be compared with those in table 1, column 3.

Broadly speaking, coefficient estimates for GDP and population have values similar to those in table 1, while the coefficient for distance continues to be estimated at around -1.3 . The coefficient for the real exchange rate (RER_{ijt}) has the expected negative sign: An increase of the real effective exchange rate that reflects a depreciation of the importing country’s currency against that of the exporting country reduces i ’s imports from j . Likewise, the remoteness variable is positive and significant: The more remote a pair of countries is from the rest of the world, the more they will tend to trade with each other. Notably, the introduction of bilateral real exchange rates and relative transport costs almost eliminates the trend for β_t , with an estimated turning point in 1981 (see figure 1). Finally, a 10 percent increase in distance would reduce bilateral

7. We thank an anonymous referee for this observation.

8. As Bergstrand (1989) points out, a negative coefficient estimate for exporter population, N_j , can be interpreted as a positive relationship between per capita income and trade (because capital-abundant countries tend to produce and export more). A negative coefficient estimate for importer population, N_i , may reflect tastes (however, this coefficient is generally not significantly different from zero, indicating an income elasticity of demand of approximately unity). Note that Coe and others (2002) introduce the population variables as a measure of geography: For larger countries, the cost of trading among themselves rather than with other countries is relatively low compared with the cost for smaller countries. This implies that large countries will tend to trade less than small countries.

TABLE 2. Distance in Standard and Augmented Panel Gravity Models

Variable	Standard		Augmented		Standard		Augmented		Standard		Augmented	
	1 All	2 All	3 P-P	4 P-P	5 R-R	6 R-R	7 P-R	8 P-R				
$\ln Y_{it}$	0.906 (44.74)	1.039 (71.33)	0.951 (27.27)	1.196 (43.65)	0.831 (17.76)	0.788 (22.53)	0.852 (14.60)	0.913 (21.21)				
$\ln Y_{it}^2$	1.164 (57.87)	1.103 (86.34)	1.352 (31.72)	1.299 (38.39)	1.005 (23.39)	0.982 (28.44)	1.303 (24.55)	1.236 (35.78)				
$\ln N_{it}$	-0.015 (0.80)	-0.072 (5.63)	0.007 (0.18)	-0.288 (8.95)	-0.020 (0.54)	-0.176 (5.76)	-0.191 (3.83)	-0.272 (7.47)				
$\ln N_{it}^2$	-0.200 (11.04)	-0.184 (15.34)	-0.350 (8.00)	-0.249 (9.15)	-0.179 (4.38)	-0.323 (10.27)	-0.298 (5.49)	-0.292 (8.59)				
$\ln D_{ij}$	-1.333 (70.46)	-1.353 (74.37)	-1.019 (39.37)	-1.059 (41.10)	-0.715 (25.62)	-0.716 (24.83)	-0.993 (38.29)	-0.978 (40.36)				
$\ln R_{it}$	0.368 (10.74)	0.525 (15.45)	0.419 (6.73)	0.692 (12.21)	0.127 (1.63)	0.085 (1.05)	0.136 (1.99)	0.061 (0.86)				
$\ln R_{it}^2$	1.909 (59.98)	2.214 (65.72)	1.647 (29.72)	1.877 (33.39)	0.165 (2.83)	0.169 (2.74)	1.658 (34.47)	1.867 (37.30)				
$\ln RER_{ijt}$	-0.0005 (6.44)	-0.0005 (6.02)	-0.0007 (5.39)	-0.0007 (5.20)	-0.0003 (2.00)	-0.0003 (2.39)	-0.0008 (4.58)	-0.0007 (7.77)				
t	-0.028 (4.76)	-0.063 (10.83)	-0.009 (1.28)	-0.038 (5.28)	-0.028 (2.64)	-0.029 (2.60)	-0.028 (3.22)	-0.0666 (4.40)				
$t \cdot \ln D_{ij}$	-0.0047 (6.80)	0.0034 (4.11)	-0.0072 (8.52)	-0.0042 (4.56)	0.0009 (1.05)	0.0048 (3.33)	-0.0069 (1.05)	-0.0053 (2.06)				
$t^2 \cdot \ln D_{ij}$	0.00012 (20.84)	0.00003 (2.46)	0.00008 (10.07)	0.0001 (8.90)	-0.00005 (3.79)	-0.00005 (3.26)	0.00007 (6.44)	0.00014 (6.44)				
$\ln K_{it}$	0.088 (6.24)	0.184 (12.18)		0.194 (12.13)		0.0727 (3.78)		0.0552 (1.96)				
$\ln K_{it}^2$				0.147 (8.35)		0.115 (5.02)		0.159 (7.51)				
$\ln P_{it}$		-0.097 (6.70)		-0.007 (0.35)		-0.408 (19.70)		-0.523 (20.77)				
$\ln \pi_{it}$		-0.216 (21.48)		-0.184 (12.60)		0.027 (1.34)		-0.153 (11.72)				
No. observations	171,998	171,998	57,332	57,332	57,332	57,332	57,332	57,332				
R-squared	0.64	0.64	0.56	0.58	0.62	0.64	0.56	0.59				
GLS vs. HT	33,548 $\chi^2(11)$	15,410 $\chi^2(15)$	900 $\chi^2(11)$	628 $\chi^2(15)$	458 $\chi^2(11)$	716 $\chi^2(15)$	804 $\chi^2(11)$	876 $\chi^2(15)$				

Source: Authors' computations based on data described in appendix A.

Note: Numbers in parentheses are t -statistics. GLS is generalized least squares estimator and HT is Hausman and Taylor (1981) estimator. P-P is bilateral trade between the poorest tercile of countries, and R-R is bilateral trade between the richest tercile.

trade by 13.4 percent in 1962 and by 13.5 percent in 1996, which is not significantly different.

However, the expected decreasing trend fails to be observed over the 35-year period. The puzzle remains. Except for explanations pertaining to the specification of the transport cost function, all possible explanations for the existence of the puzzle that have been raised in the literature have now been exhausted.⁹

Several new factors are added to the standard specification of the transportation cost function (equation 5). First, as in Limao and Venables (2001), an index is introduced for the quality of infrastructure in period t , $K_{i(j)t}$, with larger values indicating better infrastructure.¹⁰ Also included is the cost of oil, P_{Ft} , arguably the main factor affecting the marginal cost of transport. Finally, differential freight costs between primary products and manufactures are considered by including the share of primary products in total exports, π_{ijts} , as a proxy.¹¹ Because of data unavailability, π_{ijt} is proxied by π_{jts} , the share of primary export products in total exports for country j regardless of destination. Thus, the augmented transport cost function becomes:¹²

$$(8) \quad \theta_{ijt} = (K_{it})^{\rho_1} (K_{jt})^{\rho_2} (P_{Ft})^{\rho_3} (\pi_{jt})^{\rho_4} (D_{ij})^{\gamma_1 + \gamma_2 \cdot t + \gamma_3 \cdot t^2}$$

with the elasticity of transport costs to distance given by equation 6, and with the following expected signs: $\rho_1 < 0$, $\rho_2 < 0$, $\rho_3 > 0$, $\rho_4 > 0$. Inserting equation 8 into equation 7 gives the augmented gravity model:

$$(9) \quad \ln(M_{ijt}) = Z_1 \psi_1 + Z_2 \psi_2 + \beta_t \cdot \ln D_{ij} + \alpha_8 \cdot \ln K_{it} + \alpha_9 \cdot \ln K_{jt} + \alpha_{10} \cdot \ln P_{Ft} \\ + \alpha_{11} \cdot \ln \pi_{jt} + \varepsilon_{ijt}$$

The expected signs in equation 9 are $\alpha_8 = (1 - \sigma)\rho_1 > 0$, $\alpha_9 = (1 - \sigma)\rho_2 > 0$, $\alpha_{10} = (1 - \sigma)\rho_3 < 0$, $\alpha_{11} = (1 - \sigma)\rho_4 < 0$. In comparing results from the augmented

9. Also addressing the distance puzzle, Coe and others (2002) estimate a nonlinear cross-section gravity model in which they consider the possibility that the error term enters additively in equation 3 instead of as a multiplicative exponential, as implicitly assumed here. They find that when estimated nonlinearly, the coefficient for the distance variable shows a decline between 1975 and 2000, whereas it presents no clear trend when estimated with the standard log-linear specification (as in equation 7). Coe and others recognize, however, that theoretical models are nonstochastic and that they have not formally tested the appropriateness of an additive rather than a multiplicative error structure, which limits the inferences that can be drawn from their comparison of the two estimation procedures.

10. The index constructed by Limao and Venables (2001) from data in Canning (1996) is used. The index is a simple average of four indexes (roads, paved roads, telephone lines, and railways) corrected for density. Appendix A describes how this index was constructed.

11. Including the mode of transport would also be desirable, but that information is not available for such a large sample.

12. The elements in this function capture several barriers to trade beyond transport cost. The phrases *trade barrier*, *transactions costs*, and *transportation costs* are used interchangeably to remind readers that this reduced form goes beyond capturing transport costs. For a similar interpretation of the values taken by the distance coefficients in the standard gravity model, see Rauch (1999).

specification with those from the standard specification, it should be noted that the standard specification excludes the four variables in the second line of equation 9.

The coefficient estimates for the augmented transport cost function in column 2 of table 2 are relatively close in value to those obtained in column 1, even though the estimated values for the remoteness variables are now larger. All the variables included in the augmented transport cost function carry the expected signs and are significant. The coefficient for the price of oil is negative and significant. Likewise, infrastructure improvement significantly increases the volume of trade (see Limao and Venables 2001). Finally, the share of primary products in total exports of j has a significant negative impact on trade, capturing the stylized fact that freight costs are greater for primary products than for manufactures (Hummels 2001a).

Thus, with the augmented trade barrier function, the puzzle disappears. The expected decreasing trend for β_t is shown in figure 1. According to the estimates in table 2, column 2, a 10 percent increase in distance would reduce bilateral trade by 13.5 percent in 1962 and by 12.0 percent in 1996—a decrease in the impact of distance of about 11.1 percent over 35 years.

Each of the transport cost function components is also introduced separately in the standard gravity equation to see how much each contributes to solving the puzzle. Oil prices solve much of the puzzle, explaining around 45 percent of the change in the elasticity of bilateral trade to distance. Infrastructure variables are also significant, explaining about 40 percent of the puzzle. Trade composition explains about 15 percent.

Next, the results are tested for sensitivity to the standard misspecification problems of endogeneity, sample selection bias, and nonvarying coefficients. To check for sensitivity to potential endogeneity for the infrastructure, remoteness, distance, and population variables, the variables (in addition of those for GDP) are instrumented according to the Hausman and Taylor (1981) method. There is no effect on the evolution of the elasticity of bilateral trade to distance. The results are unaffected when the other coefficients are allowed to change over subperiods of three years (see appendix C, equation C4). The declining trend in the distance coefficient remains unaffected when the three variables used earlier to correct for a potential selection bias are introduced.

Finally, the dummy variables for regional agreements added in equation 9 have no effect on the estimated values for β_t (even if most of regional agreements considered have a positive and significant direct impact on bilateral trade).¹³

The conclusion is thus that correction of some of the obvious misspecifications in the gravity equation that relies on distance as a proxy for transport or

13. Trade agreements taken into account are the European Union, Mercosur, Association of Southeast Asian Nations, Andean Community, West African Economic and Monetary Union, Central African Economic and Monetary Community, Economic Community of West African States, Southern African Development Community, and Common Market for Eastern and Southern Africa.

trade barriers yields plausible results for the coefficients and also yields a plausible reduction in the barrier to trade represented by distance.

III. ROBUSTNESS

The gravity model is known to be sensitive to sample and product selection. Evenett and Keller (2002) find that it performs better than the Heckscher-Ohlin model in explaining trade patterns for manufactures, which are largely differentiated products. It has also been found repeatedly that the standard gravity model derived under the hypothesis of complete specialization in different products performs better for industrial countries than for developing economies (see Feenstra 2003, chap. 5). The suitability of the aggregate gravity model as a description of trade patterns and an approximation of barriers to trade in a heterogeneous sample of countries deserves to be explored further.

The sample is broken down into three groups of equal size with selection according to the income per capita of each bilateral trade partner so that P-P is bilateral trade between the poorest tercile of countries in each time period and R-R between the richest tercile.¹⁴

With the standard specification the overall fit is approximately the same for both groups, though coefficient values for the distance and remoteness variables are much larger for P-P bilateral trade than for R-R bilateral trade (table 2, columns 3 and 5). In the standard specification the puzzle remains for both groups, although more pronounced for the P-P group.

Results from the augmented specification for the P-P and R-R groups (columns 4 and 6 of table 2) reveal larger discrepancies across coefficients than those for the standard specification (columns 3 and 5), but larger values are again found for the coefficients for distance and remoteness for the P-P group. The coefficients for infrastructure are significant in both groups but are larger for the P-P group. The price of oil is significant only for R-R trade, whereas the share of primary commodities in trade is significant only for P-P trade.¹⁵ Thus the results from the augmented gravity model suggest the death of distance for trade between high-income countries and marginalization for low-income countries.

14. As noted by an anonymous referee, this means that the sample is likely to change depending on relative income growth and then to be endogenous to trade. However, over the whole period the maximum changes noted in the P-P group of 45 countries are the entry of 4 countries and the exit of 2. For the R-R group of 39 countries, 2 countries entered and 1 exited. These changes represent only 4.7 percent of the P-P sample observations (2740/57,332) and 3 percent of the R-R sample observations (1730/57,332).

15. Again, the introduction of variables checking for selection bias does not affect the evolution of β_r . However, the coefficient values for these variables are larger for the P-P regressions, which would be consistent with some remaining specification problems.

The results for the R-R group accord with those obtained by Hummels (2001a) for freight rate estimates for U.S. imports at the commodity level for 1974–98. Estimating freight rate costs as a function of weight, distance, commodity fixed effects, and a time trend (and a time trend squared), Hummels finds that the distance coefficient falls over time, but only after containers are introduced (in 1980). Of course, this is only very indirect evidence that an augmented trade barrier function in a gravity equation may capture some of the determinants of transport costs isolated in a more reliable data set, but it is reassuring, nonetheless.

The apparent marginalization of low-income countries is confirmed in another breakdown reporting exports from the poorest tercile of countries to the richest tercile of countries (P-R) (see table 2, columns 7 and 8). As expected, coefficients for exporters are close to the corresponding values in the P-P group and coefficients for importers are close to the corresponding values in the R-R group. The distance trend exhibits an evolution similar to that in the P-P group. In sum, the marginalization of poor countries with respect to trading partners still holds.

That said, the results, especially those for the P-P group, must be interpreted with caution. First, the diverging evolution of β_t between the standard gravity model (see table 2, columns 3 and 5) and the augmented model (columns 4 and 6) in the two samples can be explained by the rate of improvement of the infrastructure index, which is twice as large for the high-income portion of the sample as for the low-income portion. The impact of infrastructure is, in principle, controlled for through the index $K_{i(j)t}$. But at least two variables that have a bearing on the impact of distance are correlated with $K_{i(j)t}$ but are not included in the model. One is time in transit, which is higher for P-P bilateral trade.¹⁶ Another is mode of transport, which has changed more for the R-R group. Hence, when $K_{i(j)t}$ is included, $|\beta_t|$ decreases significantly for R-R while remaining unaffected in the P-P sample.

Second, explaining the puzzle for the P-P group probably requires a reconsideration of the specification of the augmented equation. One candidate is bilateral foreign direct investment, which has increased more rapidly for the R-R group and could be correlated with any of the factors independent of distance included in the model. Exchange rate volatility and various omitted trade costs could also be involved. Rose and Van Wincoop (2001) use a gravity model framework to argue that countries with currency unions trade more than three times as much with each other as with other countries. Likewise, Obstfeld and Rogoff (2001, p. 9) find that currency conversion costs and exchange rate uncertainty can boost trade costs. The same applies for informational costs associated with international trade, such as search costs (Rauch 1999), for corruption and imperfect contract enforcement (Anderson and Marcouiller

16. Using shipments of manufactures to the United States, Hummels (2001b) estimates the cost of an extra day in transit at 0.5 percent of the value shipped. At equal distance time in transit is higher for P-P bilateral trade, in part because ships travel routes less frequently.

2002), or for technological changes in transport during the period under analysis (such as containerization, which raises the quality of shipping and lags behind in poorer countries; see Hummels 2001a). Finally, the changing composition of trade over time should be investigated beyond the rough decomposition into primary and manufactured products used here.¹⁷

A final issue is the realism of the standard multiplicative form assumed for “trade frictions” in all gravity models, including this one. In equation 6 the elasticity of transport costs to distance depends only on time. Suppose instead that transport costs have two components, one that is fixed with respect to distance (such as the quality of infrastructure) and one that is variable (such as the price of oil). It can easily be shown that the elasticity of transport costs to distance could increase if the fixed cost component were falling sufficiently faster than the variable cost component (appendix D). As detailed in appendix D, efforts to estimate the resulting highly nonlinear model in the transport cost variables were unsuccessful, preventing exploration of alternatives to the standard multiplicative form for the transport cost function. However, Limao and Venables (2001) conclude that the multiplicative form of their transport cost function fit their data better than did the additive form.

IV. CONCLUSION

Several variants of a panel gravity model were used to address the distance puzzle for a sample of 130 countries over the period 1962–96. The puzzle proved robust to several ad hoc versions of the gravity model, but it was significantly reduced when the gravity model was correctly specified to include remoteness (or an index of multilateral trade resistance). Adding an augmented trade barrier function (real price of oil, index of infrastructure, and share of primary exports in total bilateral trade) that corrects for the misspecification inherent in the standard representation of transport costs by distance yielded plausible estimates of the expected death of distance.

Despite the many shortcomings associated with gravity-based indirect estimates of transport costs, several intuitively plausible results emerge from the model estimations: an elasticity of trade to income close to unity (as suggested by theory), a significant impact of the real exchange rate on the volume of bilateral trade, and expected significant signs for exporter and importer country characteristics and for the impact of remoteness on the volume of trade.

The model produces an estimate of the elasticity of trade with respect to distance that is very close to direct estimates obtained from transport cost data, and the results are consistent with those obtained using more reliable data on U.S. transport costs. The model with the augmented trade barrier function

17. Actually, changing the composition of trade toward more distance-sensitive goods can be expected to increase the elasticity of trade to distance. However, Berthelon and Freund (2004) find no support for this argument.

yields a plausible estimate of an 11 percent decrease in the impact of distance on bilateral trade over the 35-year period.

Splitting the sample into three equal-size groups by income per capita revealed significant differences in bilateral trade coefficient estimates for low-income bilateral trade and high-income bilateral trade. The coefficients capturing barriers to trade, including distance, have much higher values for the low-income group (P-P). The puzzle remains for low-income bilateral trade in the standard and the augmented models, while it remains for high-income bilateral trade only in the standard model. Even though problems of interpretation persist, the results from this sample-splitting procedure are consistent with recent claims that poor countries have been marginalized by the current wave of globalization while rich countries have benefited from a death of distance.

APPENDIX A. DATA SOURCES AND DATA PREPARATION

- M_{ijt} : Total bilateral imports, in current U.S. dollars, by country i from country j at date t . UN Comtrade. The original database does not contain any zero entries.
- $Y_{i(j)t}$: GDP of country i (j) at date t , in constant 1995 U.S. dollars. World Bank (1999).
- $N_{i(j)t}$: Total population of country i (j) at date t . World Bank (1999).
- D_{ij} : Distance in kilometres between the main city in country i and the main city in country j . Database developed by CVN. Usually, the main city is the capital city, but for some countries the main economic city is considered. The distance used is orthodromic—it takes into account the sphericity of the Earth.
- B_{ij} : Dummy variable equals 1 if i and j share a common land border, and 0 otherwise.
- $L_{i(j)}$: Dummy variable equals 1 if i (j) is a landlocked country, and 0 otherwise.
- $K_{i(j)t}$: Infrastructure index, built using four variables from the database constructed by Canning (1996): number of kilometers of roads, paved roads, railways, and number of telephone sets or lines per capita. The first three variables are ratios to the surface area (World Bank 1999) to obtain a density. Each variable is normalized with a mean equal to one. An arithmetic average is then calculated over the four variables. Data for 1996 are extrapolated because 1995 is the final year for the database.
- P_{Ft} : World oil price index. International Monetary Fund (IMF) International Financial Statistics database.
- π_{jt} : Ratio of primary export products to total exports of the country j at date t . UN Comtrade.
- RER_{ijt} : Bilateral real exchange rate. IMF International Financial Statistics database. It is computed as follows:

$$RER_{ijt} = (CPI_{jt}) / (CPI_{it}) \cdot (NER_{it/\$t} / NER_{jt/\$t}),$$

where $NER_{i(j)US,t}$ is country i 's (j 's) currency value for US\$1 at date t and $CPI_{i(j)t}$ is the consumption price index for country i (j) at date t . If the CPI is not available, the GDP deflator is used. For each pair of countries, the real exchange rate is specified such that its mean over the period is zero.

- $R_{i(j)t}$: Remoteness index defined as the weighted distance to all trading partners of country i :

$$R_{it} = \sum_j w_{jt} D_{ij} \text{ for } i \neq j \text{ and with } w_{jt} = Y_{jt} / \sum_j Y_{jt}.$$

APPENDIX B. ESTIMATION METHOD

Write the model as:

$$(B1) \quad M_{ijt} = X_{ijt}\varphi + W_{ij}\delta + \varepsilon_{ijt} \text{ with } \varepsilon_{ijt} = \mu_{ij} + \nu_{ijt},$$

where X represents k variables varying over time, and W represents g variables that are time invariant.

Some explanatory variables, such as GDP, are likely to be correlated with the bilateral specific effects, which is confirmed by the χ^2 value for the Hausman test in column 2 of table 1. This test rejects the null hypothesis of no correlation between the bilateral specific effects and the explanatory variables. Hence the generalized least squares (GLS) estimator in columns 1 and 2 of table 1 gives inconsistent estimates. The instrumental variables estimator proposed by Hausman and Taylor (1981) is used to deal with this issue.

Assume that the X_1 terms (dimension k_1) are exogenous variables, and the X_2 terms (dimension $k - k_1$) are endogenous variables (correlated with the random specific effects). Variable X_2 includes the income variables, Y_{it} and Y_{jt} . Breusch and others (1989) suggest using $[QX_1, QX_2, PX_1, W]$ as instruments, which are then taken within the model. (Here Q is a matrix that expresses deviations from country-pair means, and P is a matrix that averages the observations across time for each country pair.) The resulting estimator is consistent but not efficient because it is not corrected for heteroscedasticity and serial correlation. Following the suggestion of Hausman and Taylor (1981), first-round estimates are used to compute the variance of the specific effects, μ_{ij} , and the variance of the error term, ν_{ijt} (see, for example, Guillotin and Sevestre 1994). The instrumental variable estimator is then applied to the following transformed equation:

$$(B2) \quad [M_{ijt} - (1 - \theta)M_{ij.}] = [X_{ijt} - (1 - \theta)X_{ij.}]\varphi + (\theta W_{ij})\delta + \{\theta\mu_{ij} + [\nu_{ijt} - (1 - \theta)\nu_{ij.}]\},$$

where

$$\theta = (\sigma^2_\nu / [T\sigma^2_\mu + \sigma^2_\nu])^{1/2}$$

The test proposed by Guillotin and Sevestre (1994) is used to compare the Hausman and Taylor (HT) estimator, β_{HT} , and the GLS estimator, β_{GLS} . The Hausman statistic is based on:

$$(B3) \quad (\beta_{GLS} - \beta_{HT})[\text{var}(\beta_{HT}) - \text{var}(\beta_{GLS})]^{-1}(\beta_{GLS} - \beta_{HT})'$$

Under the null hypothesis, this test statistic is distributed as a χ^2 with degrees of freedom equal to the dimension of the vector β_{GLS} ($k + g$), constant excluded. If the calculated statistic is greater than the critical value, the null hypothesis is rejected and the Hausman and Taylor estimator is preferred to the GLS estimator. The values of the χ^2 statistic for that test in table 1 turn out to be always superior to the critical value, so that the null hypothesis is rejected and the Hausman and Taylor estimator in column 3 of table 1 is preferred to the GLS estimator when GDP variables are instrumented.

APPENDIX C. EVOLUTION OF $|\beta|$ BY SUBPERIOD

$$(C1) \quad \ln(M_{ij}) = \alpha_1 + \lambda_i + \kappa_j + \alpha_1 \cdot \ln(Y_i Y_j) + \alpha_2 \cdot \ln(N_i N_j) \\ + \beta \ln(D_{ij}) + \alpha_3 \cdot B_{ij} + \nu_{ij}$$

with $B_{ij} = 1$ if i and j share a common land border. Estimated for each year with country fixed effects (corresponding to equation 1).

$$(C2) \quad \ln(M_{ijt}) = Z_1 \psi_1 + \beta \cdot \ln(D_{ij}) + \mu_{ij} + \nu_{ijt}$$

Estimated by subperiods of three years with bilateral random effects (corresponding to equation 2).

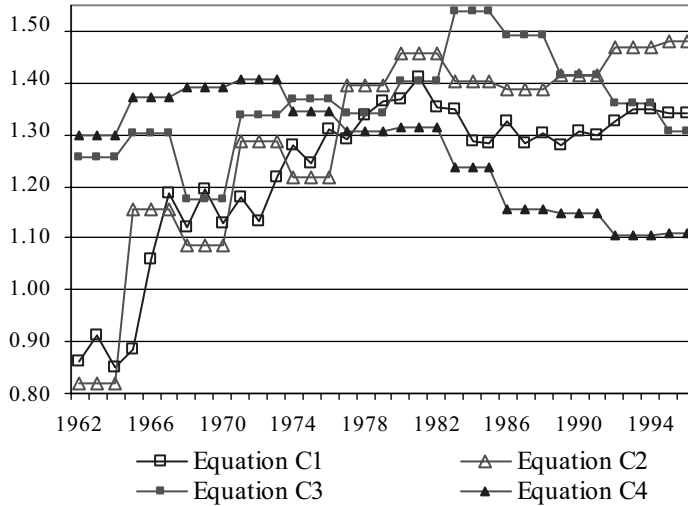
$$(C3) \quad \ln(M_{ijt}) = Z_1 \psi_1 + Z_2 \psi_2 + \beta \cdot \ln(D_{ij}) + \mu_{ij} + \varepsilon_{ijt}$$

Estimated by subperiods of three years with bilateral random effects (corresponding to equation 7).

$$(C4) \quad \ln(M_{ijt}) = Z_1 \psi_1 + Z_2 \psi_2 + \beta \cdot \ln(D_{ij}) + \alpha_8 \cdot \ln(K_{it}) \\ + \alpha_9 \cdot \ln(K_{jt}) + \alpha_{10} \cdot \ln(P_{Ft}) + \alpha_{11} \cdot \ln(\pi_{jt}) + \mu_{ij} + \nu_{ijt}$$

Estimated by subperiods of three years with bilateral random effects (corresponding to equation 9).

FIGURE C1. Evolution of the Coefficient of Distance $|\beta|$ by Subperiod



Source: Authors' computations based on data described in appendix A.

APPENDIX D. ADDITIVE TRANSPORT COST FUNCTION

Under the standard multiplicative form used in the article, the elasticity of trade costs to distance depends only on time (equation 6). Suppose now that transport costs have two components, one that is fixed with respect to distance, θ^F_{ijt} , and one that is variable, θ^V_{ijt} , that is, $\theta_{ijt} = \theta^F_{ijt} + \theta^V_{ijt}$.

In that case equation 8 would be rewritten as:

$$(D1) \quad \theta_{ijt} = \underbrace{(K_{it})^{p_1} (K_{jt})^{p_2}}_{\theta^F_{ijt}} + \underbrace{(P_{Ft})^{p_3} (\pi_{jt})^{p_4} (D_{ij})^{\gamma_1 + \gamma_2 \cdot t + \gamma_3 \cdot t^2}}_{\theta^V_{ijt}}$$

Under the specification in equation D1, it can be shown that equation 6 is now given by:

$$(D2) \quad ([\partial\theta_{ijt}/\theta_{ij}]/[\partial D_{ij}/D_{ij}]) = [\theta^V_{ijt} \times (\gamma_1 + \gamma_2 \cdot t + \gamma_3 \cdot t^2)] / (\theta^F_{ijt} + \theta^V_{ijt}).$$

From equation D2 it can be seen that the elasticity of transport costs with respect to distance could increase if the fixed cost component were falling sufficiently faster than the variable component. Substituting equation D1 into the modified version of equation 3 yields:

$$(D3) \quad \ln(M_{ijt}) = Z_1 \psi_1 + Z_2 \psi_2 + \phi \ln[(K_{it})^{p_1} (K_{jt})^{p_2} + (P_{Ft})^{p_3} (\pi_{jt})^{p_4} (D_{ij})^{\gamma_1 + \gamma_2 \cdot t + \gamma_3 \cdot t^2}]$$

where $\phi = (1 - \sigma)$. Expression D3 is highly nonlinear. Attempts to estimate equation D3 failed to yield convergence in the estimates even when estimated only for the R-R group (where errors in the variables measurement problems are likely to be smallest). The only consolation is that Limao and Venables (2001), who have real transport cost data for a sample similar to the one used here, claim that the multiplicative form is supported by their data.

REFERENCES

- Anderson, J. E., and D. Marcouiller. 2002. "Insecurity and the Pattern of Trade: An Empirical Investigation." *Review of Economics and Statistics* 84(2):342–52.
- Anderson, J. E., and E. Van Wincoop. 2003. "Gravity with Gravitas: A Solution to the Border Puzzle." *American Economic Review* 93(1):170–92.
- . 2004. "Trade Costs." *Journal of Economic Literature* 42(September):691–751.
- Baier, S. L., and J. H. Bergstrand. 2001. "The Growth of World Trade: Tariffs, Transport Costs, and Income Similarity." *Journal of International Economics* 53(1):1–27.
- Bayoumi, T., and B. Eichengreen. 1997. "Is Regionalism Simply a Diversion? Evidence from the Evolution of the EC and EFTA." In T. Ito and A. Krueger, eds., *Regionalism versus Multilateral Trade Arrangements*. National Bureau of Economic Research East Asia Seminar on Economics, vol. 6. Chicago, Ill.: University of Chicago Press.
- Bergstrand, J. H. 1989. "The Generalized Gravity Equation, Monopolistic Competition, and the Factor-Proportions Theory of International Trade." *Review of Economics and Statistics* 71(1):143–53.
- Berthelon, M., and C. Freund. 2004. "On the Conservation of Distance in International Trade." World Bank Working Paper 3293. Washington, D.C.
- Breusch, T., G. Mizon, and P. Schimdt. 1989. "Efficient Estimation Using Panel Data." *Econometrica* 57(3):695–700.
- Brun, J. F., P. Guillaumont, and J. de Melo. 1999. "La distance abolie? Critères et mesures de la mondialisation du commerce extérieur." In A. Bouet and J. Le Cacheux, eds., *Globalisation et politiques économiques: les marges de manœuvre*. Paris: Economica.
- Brun, J. F., C. Carrere, P. Guillaumont, and J. de Melo. 2002. "Has Distance Died? Evidence from a Panel Gravity Model." CEPR Discussion Paper 3500. Centre for Economic Policy Research, London.
- Canning, D. 1996. "A Database of World Infrastructure Stocks, 1950–1995." *World Bank Economic Review* 12(3):529–47.
- Cheng, I. H., and H. J. Wall. 1999. "Controlling for Heterogeneity in Gravity Models of Trade." Federal Reserve Bank of St. Louis Working Paper 99-010A. St. Louis, Mo.
- Coe, D. T., A. Subramanian, N. T. Tamirisa, and R. Bhavnani. 2002. "The Missing Globalization Puzzle." IMF Working Paper WP/02/171. International Monetary Fund, Washington, D.C.
- Deardorff, A. 1998. "Determinants of Bilateral Trade: Does Gravity Work in a Neoclassical World?" In J. A. Frankel, ed., *The Regionalization of the World Economy*. Chicago, Ill.: University of Chicago Press.
- Egger, P., and M. Pfaffermayr. 2003. "The Proper Panel Econometric Specification for the Gravity Equation: A Three-way Model with Bilateral Interaction Effects." *Empirical Economics* 28(3):571–80.
- Estevadeordal, A., B. Frantz, and A. M. Taylor. 2003. "The Rise and Fall of World Trade, 1870–1939." *Quarterly Journal of Economics* 118(May):359–407.
- Evenett, S. J., and W. Keller. 2002. "On Theories Explaining the Success of the Gravity Equation." *Journal of Political Economy* 110 (2): 281–316.
- Feenstra, R. 2003. "Increasing Returns and the Gravity Equation." *Advanced International Trade: Theory and Evidence*. Princeton, N.J.: Princeton University Press.

- Frankel, J. 1997. *Regional Trading Blocs in the World Economic System*. Washington, D.C.: Institute for International Economics.
- Guillotin, Y., and P. Sevestre. 1994. "Estimations de fonctions de gains sur données de panel: endogénéité du capital humain et effets de la sélection." *Economies et Prévision* 116(5):119–35.
- Hausman, J. A., and E. Taylor. 1981. "Panel Data and Unobservable Individual Effects." *Econometrica* 49 (6):1377–98.
- Hummels, D. 1999. "Towards a Geography of Transport Costs." University of Chicago, Graduate School of Business, Chicago, Ill.
- . 2001a. "Have International Transportation Costs Declined?" *Journal of International Economics* 54(1):75–96.
- . 2001b. "Time as Trade Barrier." Purdue University, School of Management, West Lafayette, Ind.
- Hummels, D., and V. Lugovsky. 2003. "Usable Data? Matched Partner Trade Statistics as a Measure of International Transportation Costs." Purdue University, Department of Economics, West Lafayette, Ind.
- IMF (International Monetary Fund). Various years. International Financial Statistics database. Washington, D.C.
- Leamer, A. 1993. "The Commodity Composition of International Trade in Manufactures: An Empirical Analysis." *Oxford Economic Papers* 26(3):350–74.
- Leamer, E., and J. Levinsohn. 1995. "International Trade: The Evidence." In G. M. Grossman and K. Rogoff, eds., *Handbook of International Economics*, vol. 3. New York: Elsevier, North-Holland.
- Levin, A., and C. F. Lin. 1993. "Unit Root Tests in Panel Data: New Results." University of California, San Diego, Department of Economics.
- Limao, N., and A. J. Venables. 2001. "Infrastructure, Geographical Disadvantage and Transport Costs." *World Bank Economic Review* 15(3):451–79.
- Matyas, L. 1997. "Proper Econometric Specification of the Gravity Model." *World Economy* 20(3):363–68.
- Nijman, T., and M. Verbeek. 1992. "Incomplete Panels and Selection Bias." In L. Matyas and P. Sevestre, eds., *The Econometrics of Panel Data*. Dordrecht, Netherlands: Kluwer.
- Obstfeld, M., and K. Rogoff. 2001. "The Six Major Puzzles in International Macroeconomics. Is There a Common Cause?" NBER Working Paper 7777. National Bureau of Economic Research, Cambridge, Mass.
- Rauch, J. E. 1999. "Networks versus Markets in International Trade." *Journal of International Economics* 48(1):7–35.
- Rose, J., and E. Van Wincoop. 2001. "National Money as a Barrier to International Trade: The Real Case for Currency Union." *American Economic Review, Papers and Proceedings* 91(2):386–90.
- Soloaga, I., and A. Winters. 2001. "How Has Regionalism in the 1990s Affected Trade?" *North American Journal of Economics and Finance* 12(1):1–29.
- World Bank. 1999. *World Development Indicators 1999*. CD-ROM. Washington, D.C.