Specialization: Pro- and Anti-Globalizing, 1990-2002*†

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Abstract

Specialization alters the incidence of manufacturing trade costs to buyers and sellers, with pro-and anti-globalizing effects on 76 countries from 1990-2002. The structural gravity model yields measures of Constructed Home Bias (the ratio of predicted local trade to predicted frictionless local trade) and the Total Factor Productivity effect of changing incidence. A bit more than half the world's countries experience declining CHB and rising TFP. The effects are big for the outliers. A novel test of structural gravity provides striking confirmation, validating both the CHB and TFP measures that rely on it here, and the large gravity literature that relies on it elsewhere.

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Specialization is revealed in this paper as a powerful force churning the world's economies from 1990 to 2002. Trade costs inferred from gravity are large and vary with distance, so the shifts in the location of production and consumption due to specialization and asymmetric aggregate growth (described in Section 3 below) must be changing 'average' trade costs. Structural gravity renders this intuition precise in the form of buyers' and sellers' incidence and Constructed Home Bias (CHB, the ratio of predicted local trade to predicted frictionless local trade) indexes calculated here for 18 manufacturing sectors and 76 countries.[1] The results reveal both pro- and anti-globalizing effects of specialization — some countries have falling CHB and others rising CHB, all despite unchanging bilateral trade costs. The changing buyers' and sellers' incidence has significant Total Factor Productivity (TFP) effects — some countries gain while others lose. The effects are big for the extremes.

More important to the large gravity literature, the paper provides a novel test of the structural gravity model. Its striking confirmation provides important validation for the incidence measures used here to construct CHB and TFP, and more generally a validation of applying the restrictions of structural gravity to the gravity model of trade. Structural gravity forces explain 93% of the variation in country-sector-time-directional fixed effects estimated from gravity trade flow equations, fixed effects that in principle might be expected to reflect many other forces.[2]

The CHB and TFP[3] indexes are built on measures of the buyers' and sellers' incidence of trade costs. More than half the world's countries experience declining CHB and rising TFP from changing incidence while the remainder of countries experience rising CHB and falling TFP. Fast growers with rising shares of world shipments tend to experience the biggest

[1]See Anderson and Yotov (2010) for comparable measures for Canadian provinces.

[2]Earlier gravity model estimation using aggregate trade flows often found GDP coefficient point estimates differing significantly from the theoretically indicated value of 1. But these findings are not legitimate tests of the structural gravity model because GDP is not the appropriate activity variable and aggregate gravity estimation suffers from aggregation bias that substantially understates trade costs, according to our estimates. Moreover, activity variables are perfectly collinear with the fixed effects that control for multilateral resistance, so the earlier estimates suffer from specification bias as well.

[3]Following the common abuse of terminology, TFP refers here to the Solow residual, not a pure technological change. The structural gravity model implies that the incidence of trade costs is equivalent to a technology friction, as we show below.

declines in CHB (e.g., China's home bias falls 52% and Ireland's 47% over the 12 years) and rises in TFP (China's TFP rises 18% and Ireland's 6%) while unfortunate parts of the former Soviet Union experience large increases in CHB (258% in Ukraine and 109% in Azerbaijan) and falls in TFP (19% and 42% respectively).[4] The correlation between CHB and TFP ranges from -0.7 to -0.8 depending on the approach used to calculate TFP.

Trade costs are treated as constant in the main account of the paper because the gravity model coefficients are close to constant in all the variants of the model estimated and discussed below, and including their inter-temporal variation makes essentially no difference to the main results. Constant gravity coefficients need not be inconsistent with falling trade cost *levels* because the structural gravity model implies that only *relative* trade costs can be identified: the bilateral pattern of trade is invariant to uniform proportional changes in bilateral trade costs (including costs of local shipment).

Section 1 reviews the structural gravity model and its implications to derive measures of buyers' and sellers' incidence of trade costs, CHB and the real output effect changing incidence of TFP in distribution. Section 2 describes the econometric specification. Section 3 describes the data. Section 4 presents the empirical results. Section 5 concludes.

1 The Structural Gravity Model

Our development follows Anderson (1979) in using the Armington assumption and CES preferences or technology to derive gravity. It treats the total supply to all destinations and total expenditure on goods from all origins as exogenous.[5]

[4]The churning among the world's economies revealed here is in sharp contrast to Anderson and Yotov (2010). Their portrayal of Canada's provinces over the same period uses the same methods we apply here. All Canadian provinces experienced decreased home bias and increased real output as Canada integrated with the USA and Mexico. The difference is that the present study covers (most) countries in a world economy with both fast (China, India) and slow (Japan) growing and specializing economies.

[5]The CES-Armington model nests inside many full general equilibrium models characterized by trade separability: two stage budgeting and iceberg trade costs $\Rightarrow$ distribution uses resources in the same proportions as production. The weights in the CES preferences or technology can be endogenized with monopolistic competition structure (Bergstrand, 1989) while sector supply is determined by Heckscher-Ohlin production structure. Eaton and Kortum (2002) derive an observationally equivalent structural gravity setup from ho-

Let X^k_{ij} denote the value of shipments at destination prices from origin i to destination j in goods class k. Let $t^k_{ij} \geq 1$ denote the variable trade cost factor on shipment of goods from i to j in class k, p^{*k}_i denote the factory gate price, hence destination prices are $p^{*k}_i t^k_{ij}$. Let E^k_j denote the expenditure at destination j on goods class k from all origins, while Y^k_i denotes the sales of goods at destination prices from i in goods class k to all destinations.

The CES demand function for either final or intermediate goods gives:

$$X^k_{ij} = (\beta^k_i p^{*k}_i t^k_{ij}/P^k_j)^{1-\sigma_k} E^k_j. \tag{1}$$

σ_k is the elasticity of substitution parameter for k and β^k_i is a share parameter. The CES price index is $P^k_j = [\sum_i (\beta^k_i p^{*k}_i t^k_{ij})^{1-\sigma_k}]^{1/(1-\sigma_k)}$. For deriving TFP implications we think of all goods as intermediate goods, hence P^k_j is the unit cost of the bundle of varieties and the expenditure E^k_j is spending bill on intermediate k in j.

Market clearance implies: $Y^k_i = \sum_j (\beta^k_i p^{*k}_i)^{1-\sigma_k} (t^k_{ij}/P^k_j)^{1-\sigma_k} E^k_j$. Define $Y^k \equiv \sum_i Y^k_i$ and $(\Pi^k_i)^{1-\sigma_k} \equiv \sum_j (t^k_{ij}/P^k_j)^{1-\sigma_k} E^k_j / Y^k$. Divide the market clearance equation by Y^k and use Π^k_i. Market clearance implies that world supply shares equal 'world' expenditure shares:

$$(\beta^k_i p^{*k}_i \Pi^k_i)^{1-\sigma_k} = Y^k_i / Y^k, \tag{2}$$

where $\beta^k_i p^{*k}_i \Pi^k_i$ is the world market quality adjusted price of variety i of good k. In (2) the CES 'world' price index of k is understood to be equal to 1 because summing (2) yields:

$$\sum_i (\beta^k_i p^{*k}_i \Pi^k_i)^{1-\sigma_k} = 1. \tag{3}$$

The left hand side of (2) is a CES cost or expenditure share for a hypothetical user or consumer located in the 'world' market. Π^k_i is the sellers' incidence of trade costs, as if all bilateral shipments bore a uniform incidence.

mogeneous goods combined with Ricardian technology and random productivity draws, a setup extended to multiple goods classes by Costinot and Komunjer (2007).

Now use (2) to substitute for $(\beta_i^k p_i^{*k})^{1-\sigma_k}$ in (1), the market clearance equation and the CES price index. This yields the structural gravity model:

$$X_{ij}^k = \frac{E_j^k Y_i^k}{Y^k}\left(\frac{t_{ij}^k}{P_j^k \Pi_i^k}\right)^{1-\sigma_k} \quad (4)$$

$$(\Pi_i^k)^{1-\sigma_k} = \sum_j \left(\frac{t_{ij}^k}{P_j^k}\right)^{1-\sigma_k} \frac{E_j^k}{Y^k} \quad (5)$$

$$(P_j^k)^{1-\sigma_k} = \sum_i \left(\frac{t_{ij}^k}{\Pi_i^k}\right)^{1-\sigma_k} \frac{Y_i^k}{Y^k}. \quad (6)$$

Since Π_i^k is the sellers' incidence, t_{ij}/Π_i^k is the bilateral buyers' incidence. P_j^k generated by (5)-(6) is called inward multilateral resistance and is also the CES price index of the demand system. P_j^k is interpreted as buyers' incidence because it is a CES index of the bilateral buyers' resistances on flows from i to j in class k, the weights being the frictionless trade equilibrium shares $\{Y_i^k/Y^k\}$. It is as if buyers at j pay a uniform markup P_j^k for the bundle of goods purchased on the world market.[6] Π_i^k is also called outward multilateral resistance (Anderson and van Wincoop, 2003).

Measures of 'market access' and 'supplier access' proposed by Redding and Venables (2004) are theoretically equivalent to outward and inward multilateral resistance in equilibrium but do not provide the incidence interpretation nor provide as clear a path to operationalizing the comparative statics of incidence. Their measure of 'market access' uses essentially the same formula as (5) while their measure of 'supplier access' uses the CES price index formula $P_j^k = [\sum_i (\beta_i^k p_i^{*k} t_{ij}^k)^{1-\sigma_k}]^{1/(1-\sigma_k)}$. In equilibrium the p_i^{*k}'s in the CES price index must in theory satisfy (2), so the measures are theoretically equivalent.[7]

[6] $\{P_j^k \Pi_i^k\}$ is the general-equilibrium-consistent aggregate analog to the one-good partial equilibrium incidence decomposition. If the actual set of bilateral trade costs is replaced by $\widetilde{t}_{ij}^k = P_j^k \Pi_i^k$, all budget constraints and market clearance conditions continue to hold, and factory gate prices and aggregate supply and expenditure shares remain constant (Anderson and van Wincoop, 2004). The bilateral shipment pattern shifts to that of the frictionless equilibrium. This hypothetical general equilibrium plays a role analogous to the hypothetical frictionless equilibrium in the one good partial equilibrium incidence analysis.

[7] In practice, the two measures diverge. Redding and Venables estimate a wage equation proportional to p_i^* in our setup in their one-goods-class Krugman model. We think their methods do not adequately capture the highly nonlinear interdependence of inward and outward multilateral resistance in (5)-(6), a

1.1 Calculating Incidence and Constructed Home Bias

The multilateral resistances in each sector k are solved for given $\{t_{ij}^k, E_j^k, Y_i^k\}$ from system (5)-(6). The system solves for $\{\Pi_i^k, P_j^k\}$ only up to a scalar for each class k, so an additional restriction from a normalization is needed.[8] Our empirical procedure is $P_{US}^k = 1, \forall k$. See Anderson and Yotov (2010) for more details. None of our main results depend on the normalization, as we will explain further as needed. Here we note that the trade flow in equation (4) is invariant to changes in the scalar λ of the preceding footnote.

(4) leads to a useful quantification of home bias that summarizes the effect of all trade costs acting to increase each country's trade with itself above the frictionless benchmark, $E_i^k Y_i^k / Y^k$.[9] Constructed Home Bias is defined by

$$CHB_i^k \equiv \left(t_{ii}/\Pi_i^k P_i^k\right)^{1-\sigma_k}. \tag{7}$$

CHB is the ratio of predicted to frictionless internal trade. Two regions i and j with the same internal trade cost $t_{ii} = t_{jj}$ may have quite different CHB's because $\Pi_i^k P_i^k \neq \Pi_j^k P_j^k$.

1.2 Trade Cost Incidence and TFP

Trade cost incidence shrinks the returns to national factors of production on both the output and intermediate input sides of value added. The average reduction in national Gross Domestic Product (GDP) due to the incidence of trade costs affects national TFP just like national technology frictions do. In contrast, *world* TFP in distribution from country i ag-

problem that also affects their comparative static experiments. In terms of our setup they estimate a price equation derived by equating the right hand side of (4) with the right hand side of (1) and solving for $(\beta_i p_i^*)^{1-\sigma} = \Pi_i^{\sigma-1} Y_i P_j^{\sigma-1} E_j / Y$. They estimate with regressors obtained as the fixed effects of the estimated aggregated version of trade equation (4). Recognizing that the market access and supplier access variables are co-determined with the wage (factory gate price in our case), they use lagged estimated fixed effects to replace the contemporaneous variables. They alternatively use two stage least squares with geographic instruments for origin and destination countries to proxy the fixed effects. None of these expedients adequately reflects the simultaneity of (5)-(6), in which every geographic exogenous variable in the system co-determines the fixed effect to be estimated simultaneously with every contemporaneous E_j, Y_i.

[8] If $\{\Pi_i^0, P_j^0\}$ is a solution then so is $\{\lambda \Pi_i^0, P_j^0/\lambda\}$.

[9] $t_{ij}^k = 1, \forall i, j$ implies $P_j^k = \Pi_i^k = 1, \forall i, j$, using (5)-(6).

gregates the sellers' and buyers' incidences across countries in an index developed here for perspective.

Suppose, plausibly, that imported (and local) manufactured goods are all inputs into final goods. The preceding structure justifies analyzing GDP while abstracting from the details of distribution. It is as if each sector k ships to a hypothetical world market receiving price $p_i^{Wk} = p_i^{*k}\Pi_i^k$ in each country i. Also, country i buys goods in sector k from all origins at a price index equal to one on the world market, paying the additional buyers' incidence P_i^k to bring the the bundle home. The GDP function for country i is the maximum value function $g(\{p_i^{Wk}/\Pi_i^k\}, \{P_i^k\}, q_i, v_i)$ where v_i is the national factor endowment vector and q_i is a vector of non-manufacturing goods prices. The combined reduction in GDP due to buyers and sellers incidence is implicitly defined as the scalar incidence G_i that delivers the same GDP as does the actual incidence vector $\{\Pi_i^k, P_i^k\}$:

$$G_i : g(\{p_i^{Wk}/G_i\}, \iota G_i, q_i, v_i) = g(\{p_i^{Wk}/\Pi_i^k\}, \{P_i^k\}, q_i, v_i), \tag{8}$$

where ι is the vector of ones.

Solving for local rates of change $\hat{G}_i$, using standard properties of GDP functions,[10]

$$\widehat{G}_i = -\sum_k w_i^k \widehat{\Pi}_i^k - \sum_k \omega_i^k \widehat{P}_i^k, \tag{9}$$

where w_i^k denotes the share of total manufacturing shipments from origin i due to good k and ω_i^k denotes the share of total expenditure in destination i on intermediate goods class k, and the circumflex (hat) denotes the percentage change in the multilateral resistance. (9) is the same formula used in TFP calculations treating $-\hat{\Pi}_i^k$ $(-\hat{P}_i^k)$ as a Hicks-neutral productivity improvement in output (input) technology. Conceptually $\hat{G}_i$ measures by how much manufacturing value added in i changed due to changed incidence of trade costs, all

[10]Derivatives with respect to output prices are equal to output supplies; derivatives with respect to input prices are equal to input demands. The Appendix goes into more technical detail behind the derivation of (9).

else equal. Call this the TFP effect of changing incidence.

We use (9) to calculate measures of change in manufacturing value added due to the changing incidence of trade costs using the changes in levels of multilateral resistances (one for each commodity class, country and year) and the shares of shipments and expenditures (one for each commodity class, country and year).

With a 12 year time span and discrete data, there are several plausible choices for applying the basic idea of (9). A fixed weight approach applies some sensible shipment and expenditure weights, in our case the shipment and expenditure weights for 1996, the middle year of the panel, and uses the 12 year percentage change in multilateral resistances in (9).

A more elaborate chain weights approach uses arithmetic averages of yearly adjacent shares for weights and calculates the yearly percentage change in G_j, adding the changes up to obtain the 12 year cumulative change. In practice, recognizing discrete changes, we integrate (9) over the annual intervals, hence use differentials of logs on both sides for the "price" variables. The cumulative effect is thus obtained by exponentiating

$$\ln[G_j(2002)/G_j(1990)] = -\sum_k \sum_{\tau=1}^{12} \widetilde{w}_j^k(\tau)\widehat{\Pi}_j^k(\tau) - \sum_k \sum_{\tau=1}^{12} \widetilde{\omega}_j^k(\tau)\widehat{P}_j^k(\tau) \tag{10}$$

Here,

$$\widetilde{w}_j^k(\tau) \equiv \frac{w_j^k(2002-\tau+1) + w_j^k(2002-\tau)}{2}$$

and analogously for the ω's. (10), after exponentiation, is a Tornqvist approximation to a Divisia index of G. The chain weights approach has significant advantages when countries experience large shifts in their cross-commodity composition of production and expenditure, as many do from 1990 to 2002.

Like CHB, G is invariant to the normalization used to calculate the P's and Π's in (9) or (10). This is because any rescaling raises all P's proportionally while lowering all Π's in the same proportion, the two effects canceling out in (9) and (10).

The TFP effect of national incidence of trade costs given by (8) stands in constrast to a global TFP measure of productivity in distribution. Sectoral TFP friction in distribution is defined by the uniform friction that preserves the value of sectoral shipments from origin i at destination prices: $\bar{t}_i^k = \sum_h t_{ih}^k y_{ih}^k / \sum_h y_{ih}^k$ where y_{ih}^k denotes the number of units of product class k received from i at destination h, $y_{ih}^k = X_{ih}^k / p_i^{*k} t_{ih}^k$. The iceberg metaphor captures the technological requirement that distribution requires resources in the same proportion as does production, so in product class k it takes $t_{ih}^k y_{ih}^k$ of origin i resources to deliver y_{ih}^k to destination h. $\bar{t}_i^k$ is a Laspeyres index of outward trade frictions facing seller i in good k. Aggregate TFP friction for country i is similarly given by $\bar{t}_i \equiv \sum_{k,h} \bar{t}_i^k p_i^{*k} y_{jh}^k / \sum_{k,h} p_i^{*k} y_{ih}^k$.

$\bar{t}_i$ measures the (inverse of the) productivity of distribution from i to the world economy as a whole. It measures sellers' incidence only under the partial equilibrium assumption that all incidence falls on the seller i, while the analogous inward trade cost measure would measure buyer's incidence only under the assumption that all incidence fell on buyers.

Our results show that in practice these differences are significant: Laspeyres TFP measures such as $\bar{t}_i$ and the TFP effect of incidence differ in magnitude and in the case of inward measures the correlation between them is low.

2 Econometric Specification

Two steps complete the econometric model. First, we approximate the unobservable bilateral trade costs t_{ij}^k with a set of observable variables most of which are now standard in the gravity literature:[11]

$$t_{ij}^{k\ 1-\sigma} = e^{\sum_{m=1}^{4} \beta_m^k \ln DIST_{ij}^m + \beta_5^k BRDR_{ij} + \beta_6^k LANG_{ij} + \beta_7^k CLNY_{ij} + \sum_{i=8}^{83} \beta_i^k SMCTRY_{ij}}. \tag{11}$$

Here, $\ln DIST_{ij}^m$ is the logarithm of bilateral distance between trading partners i and j. We follow Eaton and Kortum (2002) to decompose distance effects into four intervals, $m \in [1, 4]$. The distance intervals, in kilometers, are: [0, 3000); [3000, 7000); [7000, 10000); [10000,

[11]See Anderson and van Wincoop (2004) for a discussion of trade costs.

maximum].[12] $BRDR_{ij}$ captures the presence of contiguous borders. $LANG_{ij}$ and $CLNY_{ij}$ account for common language and colonial ties, respectively. Finally, $SMCTRY_{ij}$ is a set of country-specific dummy variables equal to 1 when $i = j$ and zero elsewhere.[13] These variables capture the effect of crossing the international border by shifting up internal trade, all else equal. Using the internal trade dummies has the advantage of being exogenous variables that pick up all the relevant forces that discriminate between internal and international trade. It also preserves comparability with the specification used by Anderson and Yotov for Canadian provincial trade.

Next, we substitute (11) for t_{ij} into (4) and then we expand the gravity equation with a multiplicative error term, ϵ_{ij}^k, to get:

$$X_{ij}^k = \frac{Y_i^k E_j^k}{Y^k} \left(\frac{t_{ij}^k}{\Pi_i^k P_j^k} \right)^{1-\sigma_k} \epsilon_{ij}^k. \tag{12}$$

Estimating (12) requires accounting for the unobservable multilateral resistance terms. Anderson and van Wincoop (2003) use a full information method incorporating (5)-(6). We use directional (source and destination), country-specific fixed effects, which is easier to implement and also produces consistent estimates. More important, fixed effects do not impose unitary elasticities on $P_j^{\sigma-1}, \Pi_i^{\sigma-1}, E_j, Y_i$ and allow for other unobservable country/product effects not contained in the theoretical gravity model. We use time-varying, directional, country-specific dummies to control for the MRs and the E's and Y's in the panel estimations. In Section 4.5 below we check for consistency of the estimated fixed effects with the calculated multilateral resistances and the output and expenditure shares. The results sug-

[12] Eaton and Kortum (2002) find that, with aggregate data, the estimate of the distance coefficient for shorter distances is larger (in absolute value) than for longer distances. This could reflect several forces. Motivated by transport costs, there could be mode choice switches (surface for short, air for long). But these switches could happen differently for different goods, which is particularly important for our disaggregated study, surface for heavy low value goods no matter how long the trip vs air for high value/weight goods even for quite short distances.

[13] It should be noted that we can only identify country-specific coefficients β_i^k in a panel setting, which has been used to obtain our main results. Lacking observations for enough degrees of freedom, we have to impose common cross country SMCTRY coefficients in our yearly estimates.

gest that the structural gravity model comes very close to accounting for the actual data generating process.

Another econometric challenge in gravity-type estimations is the presence of a large number of zero bilateral trade flows, which cannot be captured by a simple log-linear OLS regression. Santos-Silva and Tenreyro (2007) show that the truncation of trade flows at zero biases the standard OLS approach. In addition, they argue that not accounting for trade data heteroskedasticity in the log-linear OLS regressions produces inconsistent coefficient estimates. To account for heteroskedasticity and to utilize the information carried by the zero trade flows, Santos-Silva and Tenreyro suggest estimating the gravity equation in multiplicative form using the Poisson pseudo-maximum-likelihood (PPML) estimator.

The PPML estimator is used to obtain the main results in this study. However, in order to test the robustness of our findings, we also experiment with log-linear OLS regressions and the two-step selection procedure of Helpman, Melitz and Rubinstein (2008, henceforth HMR). HMR model selection where exporters must absorb some fixed costs to enter a market. Their model controls simultaneously for unobserved heterogeneity (the proportion of exporting firms) and for sample selection. Fixed costs provide an economic explanation, given CES preferences, for the many zeroes found in bilateral trade flows, a feature even more prominent in disaggregated data.[14] Our benchmark model is PPML because the assumptions of the HMR model, especially the exclusion restriction that permits identification, are controversial, while evidence on 10-digit bilateral US exports presented in Besedes and Prusa (2006a,b) shows flickering on and off that is hard to explain if fixed costs are important.

It is encouraging that our main results are robust to the method of estimation.It turns out that switching methods shifts the gravity coefficients equiproportionately, so that the implied t's are shifted equiproportionately (PPML and HMR reduce gravity coefficients in

[14]Our data contain 38% zeroes in 1990 and 30% zeroes in 2002. The number of zeroes varies across sectors. Naturally, the most pronounced resource sector, Petroleum and Coal, has the most zeroes, 54% in 1990 and 47% in 2002. Furniture and Beverage and Tobacco follow with about 50% zeroes in 1990 and about 40% zeroes in 2002. Santos-Silva and Tenreyro (2011) show from Monte Carlo simulations that PPML performs well even when the proportion of zeroes is large and the proportion varies with the regressors.

absolute value relative to OLS). Since gravity can only identify relative trade costs anyway, differences in estimation methods wash out with the nomalizations.

The OLS technique estimates the following econometric specification for each class of commodities in our sample:

$$\begin{aligned}\ln X_{ij,t} \quad = \quad & \beta_0 + \sum_{m=1}^{4} \beta_m^k \ln DIST_{ij,t}^m + \beta_5^k BRDR_{ij,t} + \beta_6^k LANG_{ij,t} + \beta_7^k CLNY_{ij,t} + \\ & + \sum_{i=8}^{84} \beta_i^k SMCTRY_{ij,t} + \eta_{i,t}^k + \theta_{j,t}^k + \epsilon_{ij,t}^k. \qquad (13)\end{aligned}$$

HMR differs by including a volume effect reflecting selection. PPML exponentiates (13). In (13) $X_{ij,t}$ is bilateral trade (in levels) between partners i and j at time t.[15] $\eta_{i,t}$ denotes the set of time-varying source-country fixed effects that control for the outward multilateral resistances along with total sales $Y_{i,t}^k$. $\theta_{j,t}^k$ denotes the fixed effects that control for the inward multilateral resistances along with total expenditures $E_{j,t}^k$.[16] The country fixed effects also reflect the effect of border barriers varying by country, sector and time.

3 Data Description

The data covers 1990-2002 for 76 trading partners. The countries are listed in Data Appendix.[17] Data availability allowed us to investigate 18 commodities aggregated on the basis of the United Nations' 3-digit International Standard Industrial Classification (ISIC) Revision 2.[18] To estimate gravity and to calculate the various trade cost indexes, we use

[15] Even though we experiment with yearly estimations, our main results are obtained from panel data. As noted in Cheng and Wall (2005), "Fixed-effects estimations are sometimes criticized when applied to data pooled over consecutive years on the grounds that dependent and independent variables cannot fully adjust in a single year's time." To avoid this critique, we use only the years 1990, 1994, 1998, and 2002. It turns out, however, that the gravity estimates are not very sensitive to the use of alternative lags.

[16] In a static setting, the structural gravity model implies that the income and the expenditure elasticities of bilateral trade flows are unity. Olivero and Yotov (2009) show that the income elasticities are not necessarily equal to one in a dynamic setting.

[17] Country coverage was predetermined mainly by the availability of sectoral level production data.

[18] The complete United Nations' 3-digit International Standard Industrial Classification consists of 28 sectors. We combine some commodity categories when it is obvious from the data that countries report sectoral output levels in either one disaggregated category or the other. Our commodity categories are: 1

industry-level data on bilateral trade flows and output, and we construct expenditures for each trading partner and each commodity class, all measured in thousands of current US dollars for the corresponding year.

Our focus on specialization as the driving force in the world economy is based on increasing specialization observed in falling correlation between national shares of world manufacturing output and expenditure across 18 3-digit manufacturing sectors and 76 countries from 1990 to 2002. Table 1 highlights the change in correlation overall and in a few key sectors. The decrease in total manufacturing correlation from 0.97 to 0.94 (see column 1) is

Table 1: Correlations: Output and Expenditure Shares

Year	(1) All Manufacturing	(2) Apparel (322)	(3) Leather (323)	(4) Food (311)
1990	0.97	0.94	0.82	1.00
2002	0.94	0.74	0.70	1.00

This table reports correlations between world output and expenditure shares for three 3-digit ISIC categories and total manufacturing in 1990 and 2002.

driven by sectors such as Apparel and Leather (see columns 2 and 3). In contrast, increased specialization has played no role for categories such as Food (see column 4).

Behind the correlation changes are shifts in key countries' production and expenditure shares. For example, between 1990 and 2002, China's share of world manufactures production rose from 3.1% to 7.6% while Japan's fell from 19.7% to 13.7%. Differential aggregate growth plays a role, but specialization within manufacturing induces heterogeneity in the comparative performance of sectoral shares. Japan's world share of Beverages and Tobacco actually rises while its share of Paper Products falls only marginally. China's global share of every sector rises but some less than double (e.g. Printing) while others (e.g., Apparel, Transport) quadruple.

Theoretical trade, shipments and expenditure data should be measured in user prices but valuation at full user prices is unobservable, so our data has measurement error, with

Food; 2 Beverage and Tobacco; 3 Textiles; 4 Apparel; 5 Leather; 6 Wood; 7 Furniture; 8 Paper; 9 Printing; 10 Chemicals; 11 Petroleum and Coal; 12 Rubber and Plastic; 13 Minerals; 14 Metals; 15 Machinery; 16 Electric; 17 Transportation; and, 18. Other. A detailed conversion table between ours and the UN 3-digit ISIC classification is available upon request.

implications discussed further below. We use the CEPII database that gives CIF valuation of bilateral exports for X_{ij}^{k}. While this is the best we can do, it leaves out the large portion of other trade costs paid by the users.[19] Shipments data do not include the costs other than CIF margins paid by users, and the expenditures data is constructed from shipments plus trade data.

In addition, we use data on bilateral distances, contiguous borders, colonial ties, common language, and industry level data on elasticity of substitution. The Data Appendix provides a detailed description of the data sources and the procedures used to construct all variables employed in our estimations and analysis.[20]

4 Empirical Results

4.1 Gravity Estimation Results

Tables 2-4 report the PPML panel results obtained with robust standard errors clustered by trading pair.[21] Overall, the disaggregated gravity model works well. The gravity estimates vary across commodities and across countries in a sensible way.

Importantly, we find that the coefficients are relatively stable over time for any given commodity category — the puzzle of the missing globalization persists even with our novel ability to include internal trade and thus estimate SMCTRY effects. To measure the movement of our gravity estimates over time, we construct the percentage changes in the yearly estimates for the period 1990-2002: $\%\Delta\widehat{\Theta} = 100 \times \frac{\widehat{\Theta}_{2002}-\widehat{\Theta}_{1990}}{\widehat{\Theta}_{1990}}$, where $\widehat{\Theta}$ is the yearly PPML gravity coefficient estimate of any of the regressors in our estimations.[22] Except for colonial ties (CLNY) and internal trade (SMCTRY) there is no evidence of intertemporal change

[19] Anderson and van Wincoop (2004) report on the large and varying internal distribution margins observable in aggregate data, but these exclude some important user costs falling on the end purchaser.

[20] Tables with summary statistics and the data set itself are available upon request from the authors.

[21] Comparison between estimates obtained with and without clustering reveal that the clustered standard errors are a bit larger. This suggests positive intracluster correlations, as expected.

[22] These indexes, the corresponding standard errors (computed with the Delta method), as well as the disaggregated gravity estimates for individual years, are available upon request.

while for the former the scant evidence is discussed below.

Distance (DIST1-DIST4). Distance is a large impediment to manufactures trade: all estimated distance coefficients are negative and statistically significant. For most commodities, we find an inverted u-shape (algebraically) relationship between distance and trade flows.[23] This non-monotonic pattern contrasts with results of Eaton and Kortum (2002), who report monotonically rising estimates of distance elasticities based on aggregate data. Distance elasticities vary greatly by sector, consistently with variation in value to weight and the physical requirements of transportation.

Common Border (BRDR). All estimated coefficients on the contiguity variable BRDR are positive, large and, in most cases, significant.[24] Contiguity effects vary significantly across commodities. For example, each of the positive contiguity effects for Wood, Metals and Electric products are more than twice those for Chemical products.

Common Language (LANG). All else equal, sharing a common official language facilitates bilateral trade for eleven of the categories in our sample, marginally so in the case of Rubber and Plastic Products. Insignificant LANG effects are found for more than one third of the sectors including Leather, Wood, Paper, Petroleum and Coal, Electrical Products, Transportation and Other Manufacturing products. The variation in the magnitude of the coefficients across commodities makes intuitive sense. Thus, for example, while the estimate on Paper is neither economically nor statistically significant, the estimate on Printing and Publishing is the largest. The explanation is that knowledge of a specific language is necessary for consumption of published products. It is also intuitive that language should not matter much in the case of Petroleum and Coal, which is the most pronounced resource sector in our sample.

[23] The negative effect of distance on bilateral trade is usually smallest for the countries that are less than 3000 kilometers away from each other. Depending on the commodity, the effects of distance are largest for country pairs in either the second interval, of [3000, 7000) kilometers, or in the third interval, of [7000, 10000) kilometers. In the case of Chemicals the estimate of the distance coefficient increases (in absolute value) with the distance interval.

[24] The two exceptions are Leather Products, with a positive but insignificant coefficient, and Beverages and Tobacco Products, for which the coefficient estimate on BRDR is neither economically nor statistically significant.

Colonial Ties (CLNY). As compared to the other gravity variables, 'colonial ties' is the regressor with least explanatory power. Tables 2-4 indicate that colonial ties increase bilateral, commodity-level trade flows only for a few product categories: Only one-third of the CLNY coefficient estimates are positive and significant, and, in most cases, marginally so.[25] Furthermore, the significant estimates are small in magnitude. Colonial ties matter most for the categories of Beverage and Tobacco, Apparel, Leather Products, and Printing and Publishing Products. Overall, our estimates suggest that the effects of colonial ties have slowly disappeared during the 90s.

Same Country (SMCTRY). International borders reduce trade, all else equal. The coefficient estimates reported in Tables 2-4 are averages across all country-specific estimates. The vast majority of the estimates of the coefficients on SMCTRY (the dummy variable capturing internal trade) are positive, large, and significant at any level. Beverages and Tobacco and Printing and Publishing are the two categories with highest domestic bias in trade, while Machinery and Transportation, are among the categories with lowest SMCTRY estimates. The high estimates for the former sectors suggest a large relative cost advantage (e.g., in advertising) to reaching local consumers while the low estimates for the latter sectors suggest that a lower relative cost advantage of reaching local users. The variation of the estimated SMCTRY coefficients makes intuitive sense at the country level too. First, less developed nations are more prone to buy domestic products. Thus, Mongolia, Iran and Tanzania are the three nations with largest average SMCTRY estimates, followed by Kenya, Senegal and several former Soviet republics including Kazakhstan, Kyrgyz Republic and Armenia. Second, the regions with lowest bias in purchase of domestic goods are the Netherlands, Germany, Belgium-Luxembourg and US.[26]

[25] Less than one third of the yearly estimates of the coefficients on CLNY are significant. For the few products for which the yearly estimates are significant, we find that the CLNY effects have decreased over time, but the decrease is not statistically significant.

[26] We estimate negative average internal trade coefficients for two regions: Hong Kong and Singapore. It should be noted though that the estimates for Transportation and Other manufacturing, where these regions seem to have comparative advantage are positive (but not significant), while the SMCTRY estimate on Machinery for Hong Kong is positive and statistically significant.

Overall, the SMCTRY effects are persistent over time.[27] The percentage changes of the yearly estimates are mostly small in magnitude and not statistically significant. There are a few categories for which we estimate a significant decrease in the domestic bias over the period 1990-2002. These commodities are Textiles, Apparel, Metals, and Electrical products. For Textiles and Apparel, the entry of China into the WTO is an obvious explanation despite the contrary effect of the MFA, none of which is modeled. Petroleum and Coal products is the only category for which we find an increase in domestic bias. Possible explanations for this finding include recent wars and political conflicts, and efforts to improve energy security.

In order to abstract from any effects due to changes in the yearly gravity estimates,[28] we choose to employ the panel PPML estimates in the calculation of the trade costs indexes below. As more justification for this procedure we construct correlations (by product) between the panel PPML estimates and their yearly PPML counterparts. All correlation indexes are very large (above 0.9) and significant.[29] This suggests that the two estimation sets (panel and yearly) can be used interchangeably in the calculation of the trade costs indexes.

As a robustness check, we compare the panel PPML estimates against two alternative econometric specifications, a log-linear OLS and a two-step HMR procedure. Santos-Silva and Tenreyro (2006) criticize the standard log-linear OLS approach to estimate gravity as inappropriate. They argue in favor of the PPML estimator and find significant biases in the OLS gravity estimates. In accordance with their findings, but using disaggregated data, we find an upward bias in the OLS estimates, especially in the coefficient estimates of distance and of colonial ties. These results are consistent in both the panel and the yearly estimations. Importantly for present purposes, there is almost perfect correlation between the OLS (both yearly and panel) estimates and the PPML gravity estimates. Remembering that the gravity system identifies only relative trade costs, the very high correlation between the PPML and

[27]To obtain yearly SMCTRY estimates we impose a common cross-country coefficient for each commodity, because a country-specific indicator trade dummy would have had only one observation that differs from zero.

[28]In a few instances, we obtain wrong-sign estimates. For example, we estimate a positive and significant effect of distance on bilateral trade for Textiles in 1990.

[29]These numbers, along with the yearly PPML estimates, are available upon request.

the OLS estimates suggests that the two sets are equally good proxies for the unobservable bilateral trade costs and can be used interchangeably for our purposes.

Finally, we compare the panel PPML estimates to estimates obtained with the two-step HMR method, where exporters must absorb some fixed costs to enter a market. HMR estimation controls simultaneously for sample selection bias and for unobserved heterogeneity bias. To keep the sample size as large as possible, we follow HMR in using religion as the exclusionary variable that permits identification. We apply the HMR cubic polynomial specification to control for the biases caused by the unobserved firm heterogeneity along with the Heckman (1979) sample selection correction to obtain both panel as well as yearly estimates. The disaggregated HMR estimates support their aggregate findings: both the selection and the unobserved heterogeneity biases are significant, which leads to exaggerated OLS estimates. But most important for our purposes is the almost perfect correlation between the panel PPML estimates and the HMR panel and yearly counterparts.[30] PPML and HMR (and OLS) estimates turn out to be essentially equivalent for the calculation of the trade cost indexes.

4.2 Multilateral Resistance Indexes

The multilateral resistances are calculated from (5)-(6) with fitted bilateral trade costs and data on expenditure and supply shares. Let $\tau_{ij}^k \equiv (t_{ij}^k)^{1-\sigma_k}$ the power transform of the bilateral trade cost factor. We construct τ_{ij}^k from the estimated gravity coefficients , the $\widehat{\beta}$'s, and the proxy variables data as:

$$\widehat{\tau}_{ij}^k = e^{\sum_{m=1}^{4} \widehat{\beta}_m^k \ln DIST_{ij}^m + \widehat{\beta}_5^k BRDR_{ij} + \widehat{\beta}_6^k LANG_{ij} + \widehat{\beta}_7^k CLNY_{ij} + \sum_{i=8}^{83} \widehat{\beta}_i^k SMCTRY_{ij}}.$$

Where needed to get the levels of multilateral resistances, $\widehat{\sigma}^k$ is the estimate of the elasticity of substitution obtained, as described in the data appendix, from the 3-digit HS indexes of

[30]Correlation coefficients and HMR estimation results (panel and yearly) are available upon request.

Broda et al. (2006).[31]

Theory implies that bilateral trade cost factors t_{ij}^k should always be greater than one. In less than one percent of the cases (mostly for the estimates of internal trade costs, $\widehat{t}_{ii}^k$) our estimates are lower than one. To preserve the variability in the bilateral trade cost estimates, we divide each of them by $\widehat{t}_{ij,min}^k$, where $\widehat{t}_{ij,min}^k$ is the smallest estimate for a given class of commodities (usually the internal estimate for a small country). This transformation is inconsequential for our analysis since the full gravity system is homogeneous of degree zero in the set of gravity implied trade costs. Estimation can only identify relative trade costs.

The power transforms of multilateral resistance are solved from the gravity system (5)-(6) using the $\widehat{\tau}$'s. Overall, we find significant variation, within reasonable bounds, in the multilateral resistances across countries for a single product, and across commodity lines for a given country. For brevity and clarity of exposition, we construct and analyze aggregate indexes across all commodities for a given nation, and across all countries for a given commodity category. We use expenditure shares as weights in the aggregation of the inward multilateral resistances and output shares in the aggregation of the outward counterparts.

Intertemporal variation is induced by changing shares with constant bilateral trade costs. To describe the evolution of the multilateral resistances over time, we follow the procedure from Anderson and Yotov (2010) who adopt a time-invariant normalization to convert current prices to base year prices. Applied to our setting, the procedure is to convert US current inward multilateral resistance (chosen for normalization) to base year US inward multilateral resistance. Initially we calculate MR's for each commodity with $P_{USA}(t) = 1$ for each year t. This yields (for each commodity) a set $\{P_i(t), \Pi_i(t)\}$ for each region i and year t. Using output and expenditure shares, respectively as weights, we aggregate commodity level multilateral resistances to form country MR's. To convert them to intertemporally comparable values, we construct an inflator variable for US, drawn from country-level personal consump-

[31] Estimates of σ can be obtained directly as coefficients on bilateral tariffs in the gravity model. However, due to unavailability of bilateral tariff data for the period of investigation, we choose to use the elasticity indexes from Broda et al. (2006)

tion expenditures (PCE) prices on durable and nondurable goods (but not services) for the period 1990-2002.[32] The inflator is equal to $\pi_{USA}(t) = PCE_{USA}(t)/PCE_{USA}(1990)$. The new set of 'time-consistent' MR's is $\{\pi_{USA}(t)P_i(t), (1/\pi_{USA}(t))\Pi_i(t)\}$. Conceptually, any country's inward MR is converted to a 1990 US dollars equivalent. For example, $P_i(t)/P_{USA}(t)$ is replaced by $P_i(t)/P_{USA}(1990)$. The scale of outward MR's is inversely related to the scaling of inward MR's due to the structure of the gravity system, so outward MR's are also interpreted as being in 1990 US dollars.

Deflated inward multilateral resistance (IMR) indexes are reported in Table 5. Column (1) reports IMR's by country for 1996, the midyear of our sample that is also representative.[33] The IMR values vary across countries, and the pattern of IMR variation makes good intuitive sense. More 'remote' nations, geographically and in terms of industry concentration and economic development, face larger buyers' incidence. Thus, developing countries (e.g. Mozambique, Senegal and Tanzania) and some former Soviet republics (e.g. Kyrgyz Republic and Azerbaijan) are consistently among the regions with highest buyers' incidence. In contrast, all developed countries are in the lower tail of the IMR distribution. The levels and range of values of IMR's in manufacturing for countries resemble those for Canada's provinces over a wider set of goods in Anderson and Yotov (2010).

The numbers in column (2) of Table 5 summarize the IMR changes over the period 1990-2002 and are constructed using the chain procedure of Section 1.2. Over time consumers in twenty eight of the countries in our sample enjoyed a moderate decrease in the inward multilateral resistances. Consumers in the Czech republic and Indonesia are the biggest winners. On the opposite side of the spectrum of IMR changes we find Russia, with an IMR increase of 34 percent, and another former Soviet republic, Azerbaijan, with an even larger increase of 55 percent. The estimates for these two nations are in accordance with the fact that they are among the countries with largest consumer price increases in the sample.[34]

[32] The PCE price index is constructed and maintained by the Bureau of Economic Analysis.

[33] These numbers are the deflated yearly average inward multilateral resistances for each country across all goods, weighted by the national expenditure share on each commodity.

[34] In principle, IMR changes are comparable to average CPI changes. However, IMR's may only loosely

IMR's vary across product lines for a single country. We aggregate across countries to portray the cross-commodity variation of the IMR's in Table 6. The numbers for each year and commodity are calculated as relative to the corresponding number for the United States. Thus, the indexes from column (1) of Table 6 suggest that the average consumer costs of Food, Wood, and Paper Products are considerably higher in the rest of the world, as compared to the US, while the average costs of Furniture, Apparel and Mineral Products are lower than in the US.

Estimates of the percentage changes in the IMR indexes, reported in column (2) of Table 6, show that, on average, the inward multilateral resistances have increased for most commodity categories (about two thirds) in our sample. The increase varies by product line. At 12%, Food and Furniture are the categories for which consumers have suffered the largest IMR increase over the period 1990-2002. Rubber and Plastic and Electrical products are two other sectors with large IMR increase. Wood and Beverage and Tobacco are two of the categories with IMR decrease of more than 10%. Petroleum and Coal is the category with largest IMR fall of 22%. This finding coincides with the stable oil prices of the 90's, followed by a fall in the early 2000's, and the fall in IMR's due to changing specialization patterns may be part of the explanation.[35]

Outward multilateral resistance (OMR) indexes for manufacturing also vary across industries for a single country and across nations for a single sector. Once again, to ease interpretation of our findings, we aggregate OMR's across commodities for each country and across countries for each commodity using commodity shipment shares as weights. The variation of the OMR's across countries is revealed in column (3) of Table 5, where we report

track variations in consumer price indexes and any differences between the CPI's and the IMR's have a number of explanations. First, our IMR indexes are based on a manufacturing sample, excluding services, agriculture and mining. Second, the inward incidence of trade costs probably falls on intermediate goods users in a way that does not show up in measured prices. Third, the production weighted IMR's are not really conceptually comparable to the consumer price indexes of final goods baskets. Next, home bias in preferences may be indicated by our results. Home bias in preferences results in attributions to 'trade costs' that cannot show up in prices. Finally, the IMR's are no doubt subject to measurement error and are based on a CES model that itself may be mis-specified.

[35] The oil price fall is also partly explained by a series of increases in OPEC quotas, increase in the Russian production, and a weakened US economy. Prices fell further immediately after the September 11 attacks.

deflated outward multilateral resistances for 1996. In column (4) of the table, we calculate the percentage change in the OMR's over the period 1990-2002 using the chain procedure of Section 1.2. Overall, the levels and range of values (7.8-31.9) are significantly higher than for Canada's provinces (3.5 to 7.6) in Anderson and Yotov (2010). This is also consistent with the relatively low aggregate OMR for Canada reported in Table 5. The explanation is that the largest trading partner for each Canadian province and territory, or for Canada as a whole, is very close geographically and trade between Canada and US is 'free'.

The pattern of OMR variation makes good sense for the most part. Three properties stand out. First, the regions with the largest sellers' incidence are more remote regions (geographically and in terms of industry concentration and national specialization) and some former Soviet republics. Interestingly, we find that the regions with lowest OMR's are also less developed economies such as Iran, Mongolia, Kenya and Tanzania. The reason is that these regions have the largest SMCTRY estimates, tending to lower their OMR's. Mexico also a low OMR, due to its large trade with US, fostered by NAFTA and free trade relations with the EU and series of Latin American countries during the period of investigation. As expected, we find the majority of developed countries in the lower halve of the OMR distribution. Second, OMR's are considerably larger than the IMR's, as in Anderson and Yotov (2010).[36]

Third, we find evidence of a decline over time in the OMR's for close to two-thirds of the nations in our sample and an increase in the seller's incidence for the rest of the countries. See column (4) of Table 5. This stands in sharp contrast to the uniform OMR fall for

[36]To explain why the outward multilateral resistance exceeds its inward counterpart, they draw intuition from two propositions. According to Proposition 1, larger supply or net import shares (defined as the difference between expenditure shares and output shares) tend to reduce multilateral resistances in a special case. Building on the intuition from Proposition 1, in Proposition 2 they show that a greater dispersion of supply shares, which is an empirical regularity, drives average outward above average inward multilateral resistance. This pattern also plausibly reflects specialization, with causation running from multilateral resistances to supply (and demand) allocations. The prediction of Proposition 1 holds up strikingly well in our OMR estimates, helping to explain some other results as well. Correlations between the OMR's and the output and net import shares at the product level, available upon request, reveal that the OMR's are significantly decreasing in output shares and decreasing in net import shares.

Canada's provinces in Anderson and Yotov (2010).[37] Essentially this reflects the empirical regularity that OMR decreases with supply shares and the zero sum property of changes in shares. Specialization in the global economy induces incidence shifts that cause gainers and losers, especially via sellers' incidence. The Canadian provinces turn out to be a specially favored case with the zero sum aspects of their interaction inflicted on the much larger outside world. Kuwait is the nation experiencing the largest OMR increase of close to 32 percent. The reason is the large increase in the incidence on the producers of Petroleum and Coal (Multilateral resistance and output changes for individual commodities are available upon request). Most of the oil producers in the world also suffered OMR increase, but the oil industry accounts for more than 85 percent of Kuwait's exports and for more than 50 percent of its GDP, extreme relative to other oil exporters.

Sellers in some republics from the former Soviet bloc (e.g., Armenia, Russia and Kazakhstan) enjoy significant OMR decreases, while producers in other Soviet nations, such as Ukraine, and economies under the heavy Soviet influence, such as Bulgaria, suffer large increases in seller's incidence. The explanation is that after the collapse of communism, during the early '90s, some nations lost their guaranteed portion of the large Soviet market to more competitive producers, while others expanded into new markets including some in the former Soviet Union.

The patterns of OMR variation across commodity categories and their change over time are revealed in columns (3)-(4) of Table 6. Column (3) reports deflated aggregates obtained with output shares used as weights. Leather, Petroleum and Coal and Other Manufacturing are always among the categories with highest outward trade costs at the country level, which translates into high aggregate indexes for these categories. Transportation Products and Food are consistently among the sectors with lowest OMR indexes. Results from column (4) of Table 6 indicate that over the period 1990-2002 OMR's have fallen for more than two-thirds of the commodities in the sample. Food is the category for which producers have

[37] It should be noted that we do estimate and aggregate OMR fall for Canada, which is consistent with the results from Anderson and Yotov (2010).

enjoyed the largest OMR fall of more than 12 percent. Beverage and Tobacco and Wood are two of the categories with largest OMR increase. Petroleum and Coal is the outlier with OMR increase of 22 percent.

Comparison between the changes in the incidence on consumers and producers at the product level reveals two interesting patterns. First, any IMR fall is inevitably accompanied by an OMR increase. And, second, the changes in the corresponding inward and outward multilateral resistances for each category are almost equal in absolute value. This finding reflects the formal property of (5)-(6) that the Π's and P's are inversely related on average. This nearly zero sum result is also captured by the real output numbers reported in column (7) of the table, discussed below.

The MR characteristics are in sharp contrast to those of Laspeyres trade cost (LTC) indexes. Share-weighted Laspeyres indexes of bilateral trade costs for each country and year are constructed as follows. The outward index is calculated as $\sum_k \widehat{T}_i^k Y_i^k / \sum_k Y_i^k$, where $\widehat{T}_i^k \equiv \sum_j \widehat{t}_{ij}^k X_{ij}^k / Y_i^k$. The inward counterpart calculation is $\sum_k \widehat{T}_j^k E_j^k / \sum_k E_j^k$, where $\widehat{T}_j^k \equiv \sum_i \widehat{t}_{ij}^k X_{ij}^k / E_j^k$. The inward and the outward LTC's (available on request) are very similar and the same is true for their evolution over time. There is a significant positive correlation between the OMR's and both outward and inward LTC's. In contrast, there is a weak correlation (sometimes negative) between inward multilateral resistance and either of the LTC measures. Despite the high correlation of the OMR's and the LTC's, the magnitudes are different. This difference has important resource allocation and TFP implications. An even more important difference between LTC's and multilateral resistances is that LTC's are flat over time. The pro- and anti-globalizing effects of specialization driving changes in the multilateral resistance indexes have a nearly zero-sum effect on the global efficiency LTC measures.

4.3 Constructed Home Bias

Constructed Home Bias (CHB) is calculated from (7) using the power transforms of estimated t's and multilateral resistances for each commodity and country. In practice t_{ii}^k is the estimated internal trade cost for country i and commodity k relative to the smallest internal trade cost for commodity k across all countries and regions. It is important to note that CHB is independent of the normalization, and of elasticity of substitution estimates because it is constructed using the $1 - \sigma_k$ power transforms of t's, Π's and P's.

Aggregated country level CHB indexes,[38] are obtained as weighted averages across commodity level CHB values. Recalling that CHB is interpreted as the predicted value of internal trade relative to the theoretical value of internal trade in a frictionless world, the aggregate CHB for country or region i should be the ratio of predicted internal total trade $\sum_k \widehat{X}_{ii}^k$ to frictionless internal total trade $\sum_k Y_i^k E_i^k / Y^k$. This can be obtained from (7) as:

$$CHB_i = \sum_k \left(\frac{t_{ii}^k}{\Pi_i^k P_i^k} \right)^{1-\sigma_k} \frac{Y_i^k E_i^k / Y^k}{\sum_k Y_i^k E_i^k / Y^k},$$

the weighted average of commodity CHB's, where the weights are equal to the frictionless internal trade shares.

Column (5) of Table 5 portrays the cross-national variation of CHB in 1996, which is representative of our findings across the whole period 1990-2002. Three properties stand out. First, most of the CHB values are large. There is a massive home bias in trade flows. Second, there is wide variability in CHB. Third, CHB is larger for the less developed and the smaller countries, and lower for the more developed and the industrialized nations. Thus, in each year, the United States (USA), Japan (JPN), Germany (DEU) and France (FRA) are consistently the four countries with the lowest constructed home bias indexes. On the opposite side of the CHB spectrum are developing and small countries such as Mongolia (MNG), Mozambique (MOZ), Tanzania (TZN) and Trinidad and Tobago (TTO),

[38] CHB indexes at the product level for each country are available upon request.

as are most of the former Soviet republics. Compared to the CHB's for Canadian provinces reported in Anderson and Yotov (2010), the range of CHB's is contracted substantially, explained because countries are more diverse in size than provinces.

The CHB changes over the period 1990-2002 are entirely due to the general equilibrium effect of changes in production and expenditure shares on multilateral resistance. A fall in CHB is not due to the usual understanding of globalization because the fitted trade costs t_{ij}'s (including the internal fitted values t_{ii}'s) are constant over time by definition, as we employ the panel (time-invariant) gravity estimates to construct them. Specialization drives pro- anti-globalization, both falling and rising CHB.[39]

Column (6) of Table 5 reports percentage changes in CHB for each country and region in our sample. From 1990 to 2002, CHB falls for about two-thirds of the countries while rising for the rest. The pattern is mostly explained because market share tends empirically to lower outward multilateral resistance. Slower than average growth implies smaller market shares that are associated in the cross section with higher outward multilateral resistance. The more developed and industrialized economies, which in the 90's are already pretty well set in their specialization patterns, do not experience large CHB changes. Thus Germany, Great Britain, and France are in the middle of the distribution of CHB changes. Notably, US experiences a moderate fall in CHB, while, in contrast, Japan experiences a large CHB increase. This is likely due to its lost decade of slow growth in the 90s.

The mixed CHB change results here stand in contrast those of Anderson and Yotov (2010), revealing falling CHB for all of Canada's provinces over a wider set of goods. The difference in results is due to differences in the drivers in each case, the changes in supply and demand shares. We make no attempt here to endogenize supply and demand shares but note that incidence would play a role along with other familiar forces of specialization and

[39]To assess the relative importance of specialization vs. changes in bilateral trade costs, we construct an alternative CHB measure, where we employ the yearly gravity estimates to calculate fitted trade costs. Comparison between the CHB changes obtained with the panel and the yearly gravity estimates reveals that they are almost identical. Thus, the CHB fall is indeed due to the general equilibrium effect of changes in production and expenditure shares on the multilateral resistances, and the usual forces of globalization have played a small role for the CHB changes during 1990-2002.

growth.

The newly liberalizing economies are at the upper tail of the distribution of CHB changes. For example, Mexico and China are among the nations experiencing large CHB falls. Some of the smaller European economies, such as Portugal and Greece, that have not gone through serious economic changes during the period 1990-2002, experience an increase in CHB. Finally, the big changes in former Soviet bloc trade had different effects on the CHB for the Soviet republics. Most of them suffered a significant CHB increase. Notably, three of the five countries with largest CHB increase are the Kyrgyz Republic, Ukraine and Azerbaijan. The fourth is Bulgaria, which was heavily influenced by the Soviet Union. In contrast, some former Soviet nations (e.g., Estonia, Latvia and Armenia) enjoyed CHB fall.

CHB's vary across product lines for a single country. The patterns of CHB variation across commodity categories and their change over time are revealed in columns (5) and (6) of Table 6. Notably, we find that product-level CHB indexes are relatively small and much more homogeneous as compared to their country-level counterparts. The explanation for the small magnitude is that larger and more developed countries, that have very low CHB values, are given higher weight (due to their size) in obtaining the aggregate CHB indexes for each category. Even though CHB variation across product lines is not large, it makes intuitive sense for the most part. For example, we estimate large home biases for Food and Beverage and Tobacco, which we believe are demand-driven, and for Petroleum and Coal and Mineral Products, which we explain with the nature of production and supply.[40] The low CHB values for the more technologically-oriented sectors (e.g. Machinery, Electrical products, Transportation and Other Manufacturing) are also intuitive and can be explained with clear world specialization patterns in these manufacturing sectors. Finally, we estimate an average CHB for only five of the commodities in our sample and decrease for the rest. The

[40]In particular, even though Petroleum and Coal and Mineral Products are natural resource sectors and one may expect low home biases for these categories, it should be remembered that our data covers trade and production of manufactured products, which often are produced domestically from crude natural resources that are imported, however our data does not cover natural resource trade flows. This also explains the relatively large gravity estimates of the SMCTRY coefficients, capturing internal trade, for Petroleum and Coal and Minerals.

CHB increase is the largest for Petroleum and Coal and Electrical Products, while leather and Furniture experience the largest fall.

4.4 The TFP Effect of Incidence

Changes in the incidence of trade costs impact manufacturing value added like TFP changes: real manufacturing output changes. The real output changes in the last column of Table 5, are constructed using the chain procedure of Section 1.2. It is important to note that our results for real output changes are independent of (i) the normalization used to calculate the multilateral resistances, and (ii) the choice of a price deflator (if any) used to intertemporally link the multilateral resistances. In contrast, the results *are* sensitive to the choice of elasticity of substitution in each sector, and the division of gains/losses between buyers and sellers *does* depend on the external price deflator.

Buyers in most countries suffered increasing incidence of trade costs (column 2), while sellers in about two thirds of the nations have enjoyed lower incidence of trade costs (column 4). The changes in real output, in column 7, suggest that more than half the world's countries experience rising real output while the remainder of countries experience falling real output. Globalization as specialization generates winners and losers, tending toward a zero sum.

The pattern of real output changes makes intuitive sense for the most part. In the cross section, and over time, sellers' incidence tends to fall with increases in output shares.[41] Thus, nations that have been heavily involved in trade (and wider economic) liberalization during the 90s, such as China and Mexico, are among the countries with largest real output gains.[42] Disaggregated real output changes (available upon request) reveal that these nations are consistently among the countries with largest real output gains at the commodity level. We also find that some less developed nations and economies in transition enjoy TFP gains as

[41] See Anderson and Yotov (2010) for more evidence and some theoretical insight.

[42] Venezuela is another example of a country that liberalized trade intensively during the 90s. Moreover, oil accounts for only one third of Venezuela's GDP, while manufacturing, which accounts for about one fifth of GDP, has been growing steadily at very high rates. These characteristics explain Venezuela's high real output benefit from the incidence of TFP in distribution.

well, which are driven mostly by decreasing average trade costs incidence on producers.

Most of the countries that suffer large real output decline during the 90s are some developed nations, but mostly economies in transition and members of the former Soviet bloc. In the case of the latter (see Russia for example) producers often enjoy significant decrease in outward multilateral resistance, however, the loss for consumers is even larger. These findings are supported at the commodity level as well. Along with the developing economies and the economies in transition, some developed nations, led by Japan, also experienced large real output loss. The explanation is that Japan suffered a lost decade of slow growth in the 90s. Most of the other developed nations are in the middle of the real output changes distribution.

For some nations, the real output effect of globalization is quite heterogeneous across sectors. For example, Brazil enjoyed the third largest real output gain (after China and Hong Kong, but in front of Italy) for leather products in the world. This is consistent with the fact that Brazil is one of the world's largest leather producers and it gained significantly from the geographical redistribution of the leather industry during the early 90s. More interestingly, Brazil is the leader in real output gains in the Petroleum and Coal industry. Brazilian consumers benefitted from falling oil prices during the 90s. Brazilian oil producers however, were some of the few in the world that actually enjoyed a fall in outward multilateral resistance. Brazil's share of world production rose due to the opening of the oil sector to private and foreign investors during the late 90's.

While globalization as specialization generates clear winners and losers at the country-level, our estimates reveal no significant real output changes for the world economy at the product-level over the period 1990-2002. See column (7) of Table 6. The most significant gains are realized in the Food, Minerals and Textiles categories. The overall real output effect for all manufacturing in the world economy is 0.3%. In principle, the shifts in output and expenditure shares could significantly raise this measure of world efficiency, as on a smaller scale Anderson and Yotov (2010) argued was the case for Canada's provinces, due to expenditure on average rising where inward trade costs are relatively low and similarly supply rising where outward trade costs are relatively low.[43]

Openness is positively associated with growth in our sample, as changes in CHB and changes in the incidence of TFP have a correlation coefficient varying between -0.7 and -0.8, depending on the approach used in the calculation of real output changes.[44]

4.5 Testing Structural Gravity

Structural gravity theory implies that the sum of the origin and destination country fixed effects $\widehat{\eta}^k_{i,t}+\widehat{\theta}^k_{j,t}$ estimated from equation (13) should be equal to the structural gravity term $\ln{[E^k_j P^{k\sigma_k-1}_j Y^k_i \Pi^{k\sigma_k-1}_i/Y^k]}$. A simple measure of the goodness of fit of structural gravity is based on the residuals

$$r^k_{ij,t}=\widehat{\eta}^k_{i,t}+\widehat{\theta}^k_{j,t}-\ln{[E^k_{j,t}P^{k\ \sigma_k-1}_{j,t}Y^k_{i,t}\Pi^{k\ \sigma_k-1}_{i,t}]}.$$

Since the $\widehat{\eta}^k_{i,t}+\widehat{\theta}^k_{j,t}$ variables are estimated as deviations from the US, the composite variable $\ln{[E^k_j P^{k\sigma_k-1}_j Y^k_i \Pi^{k\sigma_k-1}_i]}$ is also constructed as a deviation from the US value. In principle differencing from the US cancels out the global scaling variable $1/Y^k_t$ that is otherwise a component of the structural gravity term. The data points are very densely clustered about zero in Figure 1 showing the kernel density estimated distribution of the r's.

The structural gravity term theoretically predicts the fixed effects. The simple R^2 of structural gravity theory is 0.9.[45] Considering the amount of sector-country-time variation

[43]Totally differentiate (5)-(6), multiply by $s_i \equiv Y_i/Y$ and $b_j \equiv E_j/Y$ respectively and sum over countries. Rearranging the results: $\sum_i(s_i+\sum_j w_{ij})\widehat{\Pi}_i+\sum_j(b_j+\sum_i w_{ij})\widehat{P}_j=\frac{1}{1-\sigma}\left(\sum_{i,j}w_{ij}\widehat{b}_j+\sum_{i,j}w_{ij}\widehat{s}_i\right)$, where $w_{ij}=s_i b_j(t_{ij}/\Pi_i P_j)^{1-\sigma}$ and $\sum_{i,j}w_{ij}=1$. In the frictionless economy, the right hand side is equal to zero because the sum of shares is equal to 1. Elsewhere, the right hand side is negative as b's tend on average to rise where the composite trade friction term is large (i.e., $t_{ij}/\Pi_i P_j$ is small), and as s's tend on average to rise where the composite trade friction term is large $(t_{ij}/\Pi_i P_j)$ is small.

[44]The correlation indexes obtained with the chain-type output numbers are larger. Furthermore, if we eliminate the outliers with CHB indexes above 1000, the correlation between the changes in CHB and the changes in the incidence of TFP varies between -0.8 and -0.85.

[45]The R^2 of structural gravity is given by $1-V(r)/V(\eta+\theta)$, where $V(x)$ denotes the variance of the

in the fixed effects and in the constructed structural gravity terms this goodness-of-fit is astonishing.[46] Our prior beliefs were far more pessimistic. The standard decomposition of variance (using the theoretical coefficient of 1 on the structural gravity components) implies that the multilateral resistance terms $\ln(\Pi_i^k P_j^k)$ account for 32.3% of the variance of $\widehat{\eta}_{i,t}^k + \widehat{\theta}_{j,t}^k$'s while the size effect terms $\ln(Y_i^k E_j^k)$ account for 57.7%. That size effects should matter is hardly a surprise based on the large atheoretic gravity literature preceding structural gravity, but the large importance of the multilateral resistance term, arising strictly from structural gravity, is striking.

The fit of structural gravity improves still more by regressing the $r_{ij,t}^k$'s on sector-time fixed effects to control for differing sector-time mean measurement error in the E's and Y's.[47] In addition, to allow for time-varying, country-specific border barriers not modeled in (13) due to lack of data, we add time-varying country fixed effects:

$$\widehat{\eta}_{i,t}^k + \widehat{\theta}_{j,t}^k - \ln\left[E_{j,t}^k {P_{j,t}^k}^{\sigma_k - 1} Y_{i,t}^k {\Pi_{i,t}^k}^{\sigma_k - 1}\right] = \alpha_0 + \psi_t^k + \phi_{i,t} + e_{ij,t}^k \tag{14}$$

The right hand side fixed effects in the estimate of (14) account for 25% of the variance. Adding the additional explained variation to that of the simple R^2, structural gravity explains 92.8% of the variation in the estimated sector-country-time fixed effects $\widehat{\eta}_{i,t}^k + \widehat{\theta}_{j,t}^k$.

To probe the robustness of this finding we look suspiciously at time and country effects. As to time, the estimated ψ_t^k's and $\phi_{i,t}$'s do not vary much over time, so the structural gravity term $\ln\left[E_j^k P_j^{k\sigma_k - 1} Y_i^k \Pi_i^{k\sigma_k - 1}\right]$ absorbs essentially all time variation in the directional country-time fixed effects. To test this further, we experiment by breaking the sector-time

random variable x. The R^2's at the sectoral level are all above 0.9 for all categories except Beverages and Tobacco, Apparel, Leather Products and Petroleum and Coal Products all of which have R^2 of about 0.8.

[46]The vast majority of the fixed effects are very precisely estimated. At the same time they vary quite a bit across countries and across commodities.

[47]The estimated $\widehat{\psi}_{j,t}^k + \widehat{\phi}_{i,t}^k$ values can be rescaled as factors that shift trade flows of the gravity equation in levels, obtained by exponentiating. These range from around 0.26 to 1.73, plausibly associated with sector fixed effects on the following reasoning. Suppose, plausibly, that the mean measurement error for each goods class k is in proportion to the observable component of global shipments. The range of the estimated country-time fixed effects values is comparable to the range of values implied by the total shipments data $Y^k/[\sum_k Y^k/N]$ where N is the number of sectors. This is $(0.07, 2.51)$ and is stable over time.

and the country-time dummies into separate year, country and product fixed effects to find that none of the year dummies are quantitatively significant, and the two year coefficients that are statistically significant (for 1991 and 1993) are only marginally so.[48]

As to country effects, results are robust to eliminating country fixed effects ($\phi_{i,t} = 0$) that might control for activity elasticities not equal to one or other country-specific effects not explained by structural gravity. The R^2 of this variant of (14) falls to 0.19. Figure 2 illustrates. The x-axis shows the $r^k_{ij,t}$'s and the y-axis shows the fitted residual values, the $\widehat{e}^k_{ij,t}$'s of (14). If neither sector-time nor country fixed effects mattered, the points would lie on the 45 degree line. The 95% confidence band of the fitted line for (14) contains the line labeled "Fitted Values No Country FEs" that plots the fitted values of (14) estimated subject to $\phi_{i,t} = 0$.[49] Even the 45 degree line is mostly contained in the confidence band, emphasizing the minor role of all the fixed effects in (14) in improving the fit of structural gravity. The data points are very densely concentrated in the center of Figure 2, as indicated by Figure 1.

As a useful exploration of the performance of structural gravity we also estimate

$$\widehat{\eta}^k_{i,t} + \widehat{\theta}^k_{j,t} = \gamma^k_{0,t} + \gamma^k_{1,t} \ln \left[E^k_{j,t} {P^k_{j,t}}^{\sigma_k - 1} Y^k_{i,t} {\Pi^k_{i,t}}^{\sigma_k - 1}\right] + \epsilon^k_{ij,t} \tag{15}$$

for each k and t, and overall. Here $\epsilon^k_{ij,t}$ is a random error term that is correlated with measurement error in $\ln [E^k_{j,t} {P^k_{j,t}}^{\sigma_k - 1} Y^k_{i,t} {\Pi^k_{i,t}}^{\sigma_k - 1}]$. The standard test for estimated $\widehat{\gamma}_1 \neq 1$ is therefore invalid because the estimator is biased. Measurement error in the structural gravity term is due to the activity variables $\{Y^k_i, E^k_j\}$ both directly and in calculation of multilateral resistance in (5)-(6), along with the usual estimation error from the gravity coefficients, both of which are correlated with estimation error in the $\widehat{\eta}^k_{i,t} + \widehat{\theta}^k_{j,t}$'s.[50]

[48] Full details are available on request.

[49] For visual clarity, the data range is truncated to exclude an outlying 1.3% of dependent variable observations. About 1/4 of the country-time fixed effects are not statistically significant, although the F-test for joint significance rejects the null hypothesis for country-specific fixed effects.

[50] Theory assumes shipments evaluated at full user prices whereas actual shipments data excludes trade costs paid by users. The exclusion cancels on average in the shares but introduces unknown error to (5)-(6).

The results of estimating (15)[51] are reported in Table 7 along with a series of alternative eclectic regressions separating the components of structural gravity. The top panel of Table 7, labeled Aggregate Fixed Effects, presents the results of (15) pooled across sectors (with sector fixed effects). Columns 1-3 report yearly estimates, the column headed Panel also pools across time and the last column drops Textiles and Apparel, where the un-modeled MFA is at work. Yearly estimates are obtained with industry fixed effects, and product-year dummies are employed in the panel estimations. Our estimated γ_1's cluster very close to 1 in each specification. The constrained-uniform estimate from column 4 is 0.97. Without Textiles and Apparel, see column 5, the coefficient rises closer to 1 and the F-statistic falls.[52] In either case the standard F-test rejects the null hypothesis $\gamma_1 = 1$, but we attribute the deviation from 1 to measurement error bias in the coefficient estimate, while the F-test is also biased toward rejecting the null because the standard errors are under-estimated by not taking into account the measurement error in the right hand side variables.

Table 7 presents additionally informative results of estimating eclectic regressions to examine the performance of separate components of structural gravity. All variants indicate that multilateral resistance accounts for substantial variation in the directional fixed effects. The second panel decomposes the structural gravity term into size and multilateral resistance. Both are significant (again using standard errors that are too small) while the beta coefficients for size (available on request) indicate size is almost 3 times larger in all samples. Measurement error bias explains the coefficient estimates falling below one, downward bias that affects the multilateral resistance coefficient far more than the size coefficient. A still more eclectic approach allows for directional country fixed effects not implied by structural

For this reason, we do not generate bootstrapped standard errors for multilateral resistance.

[51] To take advantage of the additional information contained in the standard errors of the country-specific, directional fixed effects, we estimate (15) using weighted ordinary least squares with weights equal to the inverse squared standard errors of the sum of the fixed effect estimates. The intuition is that more precise estimates should be given higher weights in the estimations.

[52] Decomposing by sectors, some γ_1's are more different from 1 while others are not statistically significantly different from 1 (again using standard errors that are too small).

gravity:

$$\hat{\theta}^k_{j,t} = \alpha^k_0 + \alpha^k_2 \ln(P^k_{j,t})^{\sigma_k - 1} + \alpha^k_3 \ln E^k_{j,t} + +v^k_{j,t} \tag{16}$$

for the destination country, and

$$\hat{\eta}^k_{i,t} = \beta^k_0 + +\beta^k_2 \ln(\Pi^k_{i,t})^{\sigma_k - 1} + \beta^k_3 \ln Y^k_{i,t} + u^k_{i,t} \tag{17}$$

for the source country. See the panels labeled Inward Effects and Outward Effects in Table 7 for the results pooled across sectors. The coefficients deviate further from 1 compared to the top panel.

The results give useful clues for amending the theory. First, some sectors are less well described by structural gravity than the others, especially those for Textiles and Apparel where the Multi-Fibre Arrangement is at work. Dropping these sectors moves γ_1 closer to 1. A better approach, beyond the scope of this study, is to specify bilateral trade costs for these sectors to capture the effect of the MFA and other important trade policies. Second, combined with imposing the structural gravity value of coefficients, the results of the eclectic regressions suggest some properties of measurement error that it may be possible to utilize in the service of more accurate estimation.

4.6 Aggregation Bias

Large aggregation bias is revealed comparing our results with gravity coefficients estimated on aggregates of the 18 industries. Distance elasticities and border effects from aggregate data are reduced by a third or more in absolute value. More important, there is downward bias in CHB. The bias is skewed toward the less developed countries and close to zero for the most developed countries.

CHB changes and real output changes from aggregate data are reduced in average absolute magnitude, implying dispersion is reduced. The good performance of the gravity model on disaggregated data is notable in relation to the earlier gravity literature, which encountered poorer performance with disaggregated estimation.[53]

5 Conclusion

Specialization acting on the incidence of trade costs is revealed in this paper to be a powerful force moving the world's economies. The results show that more than half of the 76 countries experience significant declines in Constructed Home Bias while the remainder experience rising CHB. Changes in buyers' and sellers' incidence raise real output of more than half the countries while lowering real output of the remainder. At the extremes, the changes in CHB and the real output effects of incidence changes are big.

The structural gravity model is revealed as remarkably successful in fitting disaggregated manufacturing data. Future work should address the impact of measured border restrictions and proxies for non-border variables that act differentially on international trade, confronting the issues of endogeneity that arise. Future research should also include development of data that would permit extending the coverage of structural gravity estimation to disaggregated trade flows in non-manufacturing sectors. Regional trade flow analysis for Canada (Anderson and Yotov, 2010) reveals large within-country differences in the implications of specialization. Future research should include improvements in regional trade flow data for other countries as an aid to understanding the sharp differences between regions that emerge with globalization.

[53]The earlier gravity literature utilized GDP's of origin and destination countries and often found elasticities that differed from 1. Previous estimates are dubious for two reasons apart from aggregation bias. First, the appropriate activity variables are for total shipments, not a value added concept like GDP. Second, gravity models without controls for multilateral resistance have omitted variable bias (Anderson and van Wincoop, 2003). Both are controlled by direction-country-time fixed effects.

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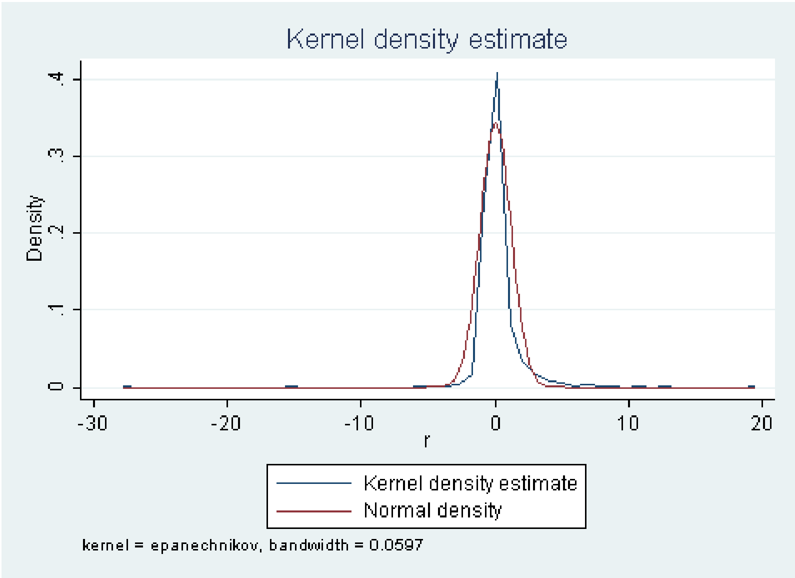

Figure 1: Estimated Distribution of r

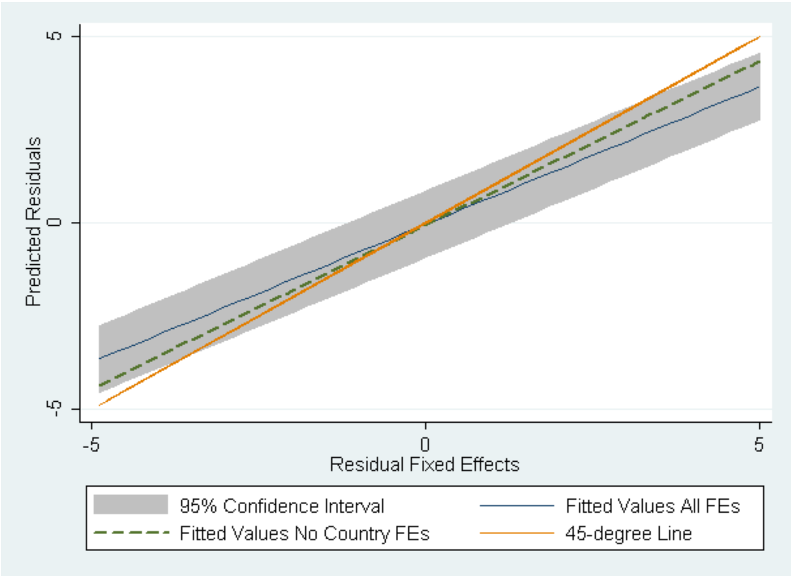

Figure 2: Residual Fixed Effects

Table 2: PPML Panel Gravity Estimates

	Food	BevTob	Textiles	Apparel	Leather	Wood
DIST1	-0.887	-0.650	-0.505	-0.578	-0.395	-1.288
	(0.117)**	(0.171)**	(0.085)**	(0.145)**	(0.199)*	(0.110)**
DIST2	-0.949	-0.776	-0.605	-0.699	-0.454	-1.308
	(0.107)**	(0.152)**	(0.080)**	(0.133)**	(0.175)**	(0.108)**
DIST3	-0.950	-0.755	-0.643	-0.683	-0.425	-1.272
	(0.101)**	(0.151)**	(0.073)**	(0.126)**	(0.162)**	(0.088)**
DIST4	-0.910	-0.730	-0.635	-0.657	-0.440	-1.283
	(0.096)**	(0.141)**	(0.071)**	(0.125)**	(0.154)**	(0.094)**
BRDR	0.505	0.028	0.516	0.522	0.323	0.724
	(0.118)**	(0.262)	(0.105)**	(0.169)**	(0.233)	(0.144)**
LANG	0.397	0.596	0.429	0.522	0.312	0.134
	(0.112)**	(0.195)**	(0.111)**	(0.178)**	(0.213)	(0.149)
CLNY	0.162	0.441	0.043	0.430	0.484	0.188
	(0.151)	(0.180)*	(0.190)	(0.182)*	(0.261)+	(0.116)
SMCTRY	4.231	5.938	3.634	3.882	4.276	3.273
	(0..275)**	(0..405)**	(0.213)**	(0.361)**	(0.436)**	(0.258)**
CONST	17.261	-9.116	13.515	10.925	10.785	18.368
	(0.995)**	(1.555)**	(0.743)**	(1.606)**	(1.496)**	(0.945)**
N	21172	21033	21172	21172	21096	21172
LL	-1.751e+08	-7.015e+07	-1.097e+08	-1.525e+08	-4.546e+07	-3.988e+07

Notes: + $p < 0.10$, * $p < .05$, ** $p < .01$. Huber-Eicker-White clustered standard errors are reported in parentheses. Time-varying, directional, country-specific fixed effects are used in each estimation. The years included in the estimations are 1990, 1994, 1998, and 2002. The estimate on SMCTRY is an average of the corresponding country-specific coefficients. Standard errors for SMCTRY are calculated with the Delta method.

Table 3: PPML Panel Gravity Estimates

	Furniture	Paper	Priniting	Chemicals	PetrCoal	RbbPlst
DIST1	-1.216	-0.964	-0.798	-0.602	-1.165	-0.741
	(0.138)**	(0.088)**	(0.121)**	(0.079)**	(0.113)**	(0.083)**
DIST2	-1.187	-1.009	-0.860	-0.648	-1.153	-0.822
	(0.116)**	(0.079)**	(0.106)**	(0.069)**	(0.101)**	(0.074)**
DIST3	-1.204	-1.045	-0.831	-0.688	-1.191	-0.830
	(0.115)**	(0.074)**	(0.104)**	(0.067)**	(0.094)**	(0.070)**
DIST4	-1.161	-1.011	-0.788	-0.689	-1.222	-0.802
	(0.111)**	(0.072)**	(0.099)**	(0.062)**	(0.092)**	(0.068)**
BRDR	0.522	0.528	0.673	0.266	0.662	0.656
	(0.130)**	(0.101)**	(0.129)**	(0.085)**	(0.121)**	(0.100)**
LANG	0.444	0.093	1.180	0.250	0.053	0.171
	(0.199)*	(0.097)	(0.138)**	(0.124)*	(0.149)	(0.102)+
CLNY	0.264	-0.136	0.365	-0.073	0.224	0.073
	(0.138)+	(0.152)	(0.141)**	(0.128)	(0.143)	(0.142)
SMCTRY	3.879	3.470	6.019	3.540	4.228	4.149
	(0.295)**	(0.209)**	(0.282)**	(0.184)**	(0.257)**	(0.200)**
CONST	18.043	18.323	11.886	15.431	17.747	14.761
	(1.434)**	(0.764)**	(1.243)**	(0.706)**	(0.991)**	(0.791)**
N	21172	21172	21172	21172	20944	21172
LL	-2.979e+07	-5.944e+07	-2.257e+07	-2.942e+08	-1.191e+08	-5.135e+07

Notes: + $p < 0.10$, * $p < .05$, ** $p < .01$. Huber-Eicker-White clustered standard errors are reported in parentheses. Time-varying, directional, country-specific fixed effects are used in each estimation. The years included in the estimations are 1990, 1994, 1998, and 2002. The estimate on SMCTRY is an average of the corresponding country-specific coefficients. Standard errors for SMCTRY are calculated with the Delta method.

Table 4: PPML Panel Gravity Estimates

	Minerals	Metals	Machinery	Electric	Transport	Other
DIST1	-1.031	-0.773	-0.433	-0.298	-0.504	-0.262
	(0.082)**	(0.073)**	(0.094)**	(0.101)**	(0.136)**	(0.119)*
DIST2	-1.039	-0.797	-0.474	-0.353	-0.593	-0.303
	(0.074)**	(0.066)**	(0.082)**	(0.090)**	(0.117)**	(0.105)**
DIST3	-1.040	-0.828	-0.506	-0.414	-0.591	-0.338
	(0.070)**	(0.063)**	(0.078)**	(0.083)**	(0.114)**	(0.097)**
DIST4	-1.039	-0.815	-0.480	-0.372	-0.575	-0.327
	(0.068)**	(0.061)**	(0.074)**	(0.082)**	(0.116)**	(0.093)**
BRDR	0.527	0.704	0.362	0.723	0.575	0.523
	(0.101)**	(0.082)**	(0.107)**	(0.136)**	(0.147)**	(0.117)**
LANG	0.230	0.288	0.258	-0.041	0.147	-0.127
	(0.080)**	(0.101)**	(0.097)**	(0.106)	(0.133)	(0.151)
CLNY	0.099	0.051	-0.062	-0.021	-0.305	0.191
	(0.088)	(0.130)	(0.110)	(0.145)	(0.253)	(0.116)+
SMCTRY	4.481	3.173	3.543	4.587	3.865	4.475
	(0.193)**	(0.178)**	(0.212)**	(0.245)**	(0.322)**	(0.259)**
CONST	16.646	17.972	13.504	10.963	14.161	11.271
	(0.702)**	(0.704)**	(0.928)**	(1.043)**	(1.270)**	(1.046)**
N	21172	21172	21172	21172	21172	21172
LL	-3.397e+07	-2.808e+08	-3.268e+08	-3.388e+08	-5.035e+08	-1.374e+08

Notes: + $p < 0.10$, * $p < .05$, ** $p < .01$. Huber-Eicker-White clustered standard errors are reported in parentheses. Time-varying, directional, country-specific fixed effects are used in each estimation. The years included in the estimations are 1990, 1994, 1998, and 2002. The estimate on SMCTRY is an average of the corresponding country-specific coefficients. Standard errors for SMCTRY are calculated with the Delta method.

Table 5: Trade Costs and Output Indexes by Country

ISO	(1) IMR_{1996}	(2) $\%\Delta IMR_{90/02}$	(3) OMR_{1996}	(4) $\%\Delta OMR_{90/02}$	(5) CHB_{1996}	(6) $\%\Delta CHB_{90/02}$	(7) Output
ARG	1.1	13.5	11.5	3.6	93	83	-17.0
ARM	1.4	3.5	14.5	-37.3	6717	-88	33.8
AUS	1.4	3.9	12.1	2.3	63	18	-6.2
AUT	1.0	1.5	13.3	3.0	49	10	-4.4
AZE	1.4	54.8	17.6	-13.1	1370	109	-41.7
BGR	0.9	12.9	18.7	17.7	437	156	-30.6
BLX	1.0	4.9	14.0	0.9	16	24	-5.8
BOL	1.4	14.0	13.0	-18.8	1627	-10	4.8
BRA	1.1	8.6	11.8	1.1	38	61	-9.7
CAN	1.1	5.9	11.1	-3.6	23	6	-2.2
CHL	1.3	5.1	12.9	-7.1	211	-11	2.0
CHN	1.0	-10.1	12.4	-7.8	13	-52	17.9
COL	1.2	4.4	12.5	-7.6	216	-20	3.2
CRI	1.4	3.6	12.8	-11.6	801	-31	8.0
CYP	1.3	1.6	14.3	-3.2	982	7	1.5
CZE	1.0	-15.0	13.2	17.1	91	-4	-2.1
DEU	0.9	-3.0	12.5	2.2	7	-0	0.8
DNK	1.1	-0.6	14.4	-1.8	63	0	2.4
ECU	1.3	4.9	15.5	-15.8	868	-38	10.9
EGY	1.3	3.3	14.3	7.0	259	51	-10.3
ESP	1.0	0.5	12.9	-1.3	27	-2	0.9
EST	1.3	-1.5	19.2	-7.2	893	-61	8.7
FIN	1.0	-0.2	14.7	4.7	84	-1	-4.5
FRA	0.9	1.3	12.3	2.0	11	0	-3.3
GBR	1.0	4.1	13.1	-0.8	13	13	-3.3
GRC	1.2	5.3	15.5	0.6	152	31	-5.8
GTM	1.3	4.6	11.0	-24.0	840	-56	19.4
HKG	1.3	0.0	18.5	-8.7	12	-11	8.6
HUN	1.1	-0.7	15.4	-1.5	200	-34	2.2
IDN	1.3	-12.2	14.3	5.5	70	-30	6.6
IND	1.1	-12.1	9.4	7.7	57	-12	4.4
IRL	1.1	-0.6	18.0	-5.0	63	-47	5.6
IRN	1.1	2.1	7.8	-12.4	223	-34	10.3
ITA	1.0	-3.1	14.7	0.4	13	-14	2.7
JOR	1.3	-2.3	11.7	-4.8	1451	-29	7.1
JPN	1.0	5.1	10.4	1.9	5	41	-7.0
KAZ	1.3	-4.3	12.1	-18.2	1004	-80	22.6
KEN	1.1	-0.4	8.6	-4.3	567	-62	4.6
KGZ	1.6	23.0	12.4	-1.7	8044	273	-21.3
KOR	1.1	-5.4	12.2	-0.8	20	-29	6.3
KWT	1.3	-8.9	31.9	31.9	140	19	-23.1
LKA	1.4	3.5	18.6	-16.4	582	-58	12.9
LTU	1.2	-11.2	19.5	-0.1	823	-50	11.2
LVA	1.3	-8.2	16.1	-13.5	1064	-77	21.7
MAR	1.2	6.7	11.9	-9.7	359	-17	3.0
MDA	1.4	-2.2	15.0	0.5	2657	-77	1.7
MEX	1.3	1.7	10.0	-13.2	56	-52	11.5
MKD	1.2	26.2	14.6	-14.0	2153	87	-12.2
MLT	1.3	1.7	20.8	-1.0	979	-2	-0.6

Continued

Table 5 – continued from previous page

ISO	(1) IMR_{1996}	(2) $\%\Delta IMR_{90/02}$	(3) OMR_{1996}	(4) $\%\Delta OMR_{90/02}$	(5) CHB_{1996}	(6) $\%\Delta CHB_{90/02}$	(7) Output
MNG	1.5	26.8	8.4	5.6	9722	324	-32.5
MOZ	1.7	-2.5	10.9	-5.5	5939	-62	8.0
MUS	1.6	-0.9	22.1	-6.7	798	-48	7.7
MYS	1.3	-7.4	16.8	-3.5	36	-50	10.9
NLD	1.0	3.5	15.2	-2.8	13	-9	-0.6
NOR	1.1	5.7	11.9	-3.3	110	14	-2.4
OMN	1.4	6.4	13.2	-20.9	902	-48	14.6
PAN	1.4	5.6	18.4	-15.2	1221	-6	9.6
PHL	1.3	2.2	13.4	-8.1	136	-28	5.9
POL	1.0	-5.1	11.0	-3.2	93	-33	8.4
PRT	1.1	5.9	16.0	-3.9	87	16	-2.0
ROM	1.0	7.6	12.3	0.5	263	-36	-8.1
RUS	1.2	33.5	13.2	-27.5	54	69	-6.0
SEN	1.6	13.0	10.7	-16.2	1736	-3	3.2
SGP	1.4	-3.8	22.0	-0.0	20	-19	3.8
SLV	1.5	1.8	15.2	-19.9	1568	-55	18.1
SVK	1.1	-2.3	16.0	2.5	324	-18	-0.2
SVN	1.1	3.7	15.0	-4.3	352	-1	0.6
SWE	1.1	2.4	13.8	3.3	51	17	-5.7
TTO	1.4	6.2	24.3	-11.7	2005	-64	5.5
TUR	1.1	0.3	12.0	-1.8	79	-10	1.6
TZA	1.5	8.1	8.7	-29.2	3274	-48	21.1
UKR	1.1	-7.3	11.1	26.5	290	258	-19.2
URY	1.3	12.1	14.8	-3.3	684	-3	-8.8
USA	1.1	7.3	12.6	-9.8	3	-11	2.5
VEN	1.2	-11.5	12.9	-15.9	280	-70	27.3
ZAF	1.2	-2.0	13.3	12.4	136	35	-10.4

Columns (1) and (3) of this table report deflated IMR and OMR levels, respectively by country for 1996. IMR (OMR) indexes are obtained from the corresponding individual commodity category values with expenditure (output) shares used as weights. IMR and OMR changes over the period 1990-2002, constructed using the chain procedure described in the text, are reported in columns (2) and (4). CHB levels and changes by country are presented in columns (5) and (6). Column (7) reports real output changes by country over the period 1990-2002.

Table 6: Trade Cost and Output Indexes by Product

Prod. Descr	(1) IMR_{1996}	(2) $\%\Delta IMR_{90/02}$	(3) OMR_{1996}	(4) $\%\Delta OMR_{90/02}$	(5) CHB_{1996}	(6) $\%\Delta CHB_{90/02}$	(7) Output
Food	1.22	11.58	5.11	-12.37	9.1	-1.3	0.79
Bev_Tob	1.02	-11.67	13.28	11.60	9.8	-5.6	0.07
Textiles	1.07	3.33	7.72	-4.50	7.4	-19.2	1.18
Apparel	0.87	-0.84	24.92	0.67	5.8	-21.5	0.17
Leather	0.98	2.85	31.50	-3.08	5.5	-42.3	0.24
Wood	1.22	-14.94	14.13	14.87	5.6	-13.1	0.07
Furniture	0.83	11.76	24.44	-12.30	6.8	-37.2	0.54
Paper	1.26	6.21	9.02	-6.21	5.5	2.1	-0.00
Printing	1.14	4.90	17.43	-4.92	5.5	-13.8	0.03
Chemicals	1.10	2.73	11.73	-2.79	6.2	-3.0	0.06
Petr_Coal	1.07	-22.47	27.03	21.90	9.5	26.8	0.57
Rbb_Plst	0.94	8.14	9.99	-8.20	6.4	-4.9	0.06
Minerals	0.88	1.48	20.46	-2.21	9.6	-1.5	0.74
Metals	1.03	-1.68	6.17	1.70	6.4	0.7	-0.03
Machinery	1.00	5.99	10.25	-5.97	4.1	-3.0	-0.02
Electric	0.95	7.28	18.40	-7.27	4.4	12.3	-0.01
Transport	1.04	6.01	3.55	-6.01	4.4	1.9	-0.00
Other	1.13	5.54	39.94	-5.64	3.5	-2.7	0.10

Columns (1) and (3) of this table report deflated IMR and OMR levels, respectively by product for 1996. IMR (OMR) indexes are obtained from the corresponding individual commodity category values with expenditure (output) shares used as weights. IMR and OMR changes over the period 1990-2002, constructed using the chain procedure described in the text, are reported in columns (2) and (4). CHB levels and changes by product are presented in columns (5) and (6). Column (7) reports real output changes by product over the period 1990-2002.

Table 7: Gravity Fixed Effects Composition

	(1) 1990	(2) 1996	(3) 2002	(4) PANEL	(5) NOTXTL
AGGREGATE FIXED EFFECTS					
$\ln Y_i^k E_j^k P_j^{k\,\sigma_k-1} \Pi_i^{k\,\sigma_k-1}$	0.950	0.973	0.970	0.969	0.978
	(0.016)**	(0.008)**	(0.009)**	(0.003)**	(0.003)**
cons	-54.061	-56.250	-56.703	-55.015	-55.507
	(0.805)**	(0.437)**	(0.488)**	(0.164)**	(0.161)**
N	1097	1350	1350	16972	15087
R^2	0.949	0.951	0.922	0.940	0.948
$F(1,\cdot)$	16.09	14.67	12.49	167.04	87.97
SIZE vs. MR EFFECTS					
$\ln Y_i^k E_j^k$	0.887	0.914	0.896	0.909	0.927
	(0.030)**	(0.013)**	(0.014)**	(0.004)**	(0.004)**
$\ln P_j^{k\,\sigma_k-1} \Pi_i^{k\,\sigma_k-1}$	0.720	0.724	0.635	0.718	0.766
	(0.074)**	(0.034)**	(0.056)**	(0.012)**	(0.012)**
cons	-49.763	-50.759	-41.261	-50.787	-51.900
	(1.822)**	(0.833)**	(0.737)**	(0.284)**	(0.282)**
N	1097	1350	1350	16972	15087
R^2	0.952	0.954	0.927	0.943	0.950
INWARD EFFECTS					
$\ln P_j^{k\,\sigma_k-1}$	0.375	0.407	0.243	0.377	0.400
	(0.030)**	(0.024)**	(0.087)**	(0.011)**	(0.012)**
$\ln E_j^k$	0.774	0.824	0.781	0.808	0.821
	(0.016)**	(0.013)**	(0.034)**	(0.006)**	(0.006)**
cons	-20.313	-21.574	-20.677	-21.108	-21.424
	(0.388)**	(0.326)**	(0.877)**	(0.140)**	(0.151)**
R^2	0.908	0.918	0.869	0.907	0.914
$\ln E_j^k P_j^{k\,\sigma_k-1}$	0.804	0.873	0.823	0.854	0.866
	(0.028)**	(0.016)**	(0.044)**	(0.007)**	(0.007)**
cons	-21.386	-23.258	-22.301	-22.577	-22.865
	(0.665)**	(0.387)**	(1.069)**	(0.164)**	(0.177)**
R^2	0.807	0.836	0.755	0.824	0.830
N	1098	1350	1350	16992	15104
OUTWARD EFFECTS					
$\ln \Pi_i^{k\,\sigma_k-1}$	0.390	0.489	0.556	0.494	0.502
	(0.072)**	(0.040)**	(0.031)**	(0.014)**	(0.014)**
$\ln Y_i^k$	0.779	0.841	0.834	0.839	0.843
	(0.024)**	(0.016)**	(0.010)**	(0.006)**	(0.006)**
cons	-22.362	-24.899	-25.330	-24.275	-24.406
	(0.868)**	(0.541)**	(0.373)**	(0.201)**	(0.206)**
R^2	0.901	0.911	0.924	0.893	0.899
$\ln Y_i^k \Pi_i^{k\,\sigma_k-1}$	0.808	0.879	0.853	0.872	0.876
	(0.025)**	(0.015)**	(0.009)**	(0.006)**	(0.006)**
cons	-25.166	-27.878	-27.494	-26.955	-27.081
	(0.704)**	(0.423)**	(0.272)**	(0.205)**	(0.207)**
R^2	0.868	0.895	0.913	0.877	0.882
N	1097	1350	1350	16972	15087

Standard errors in parentheses. + $p < 0.10$, * $p < .05$, ** $p < .01$. The yearly estimates are obtained with product fixed effects. The panel estimates are obtained with product-year fixed effects. Fixed effects estimates are omitted for brevity. The estimator is weighted least squares. See text for details.

Appendix A: Technical Appendix

Normalization

Normalization must ultimately be consistent with both the gravity module and the general equilibrium superstructure. Note that the CES price index is given by

$$P_j^{1-\sigma} = \sum_i (\beta_i p_i^*)^{1-\sigma} t_{ij}^{1-\sigma}. \tag{18}$$

Imposing $P_j = 1$ implies, for given (estimated) t_{ij}'s, a normalization on the unknown $\beta_i p_i^*$'s. Then $(\beta_i p_i^*)^{1-\sigma}$ can be solved for using the CES demand for each good i in location j:

$$X_{ij}/E_j = (\beta_i p_i^*)^{1-\sigma} t_{ij}^{1-\sigma}, \forall i.$$

The resulting $(\beta_i p_i^*)^{1-\sigma}$'s can in principle be used to check that the theoretical normalization of the Π's is met by the Π's calculated with the actual procedure: set $P_j = 1$ and drop the jth CES price index equation $P_j^{1-\sigma} = \sum_i (t_{ij}/\Pi_i)^{1-\sigma} Y_i/Y$ from the MR system. Because the multilateral resistance system is constructed and solved using $Y_i/Y = (\beta_i p_i^* \Pi_i t_{ij})^{1-\sigma}$ it is automatically consistent with the theoretical normalization.

Some other choice such as $P_{j+1} = 1$ will change the solution to multilateral resistances but with no implications for allocation. Demand system estimation can only ever identify the β's up to a scalar, so picking some $P_j = 1$ is effectively normalizing the estimated coefficients $\{\beta_i\}$.

TFP Effect of Incidence

Strictly speaking, (9) does not follow from differentiating (8) unless manufacturing output and input demand are separable in the GDP function. Generally, the first (second) element on the right hand side of (9) should be multiplied by the ratio of manufacturing sales (purchases) in the actual equilibrium to manufacturing sales (purchases) in the hypothetical equilibrium.

The simpler expression (9) is restored if G_i is redefined for local comparison purposes. Consider a scalar measure comparing two adjacent years 0 and 1 with prices $\{p_i^{Wk}(0), P_i^k(0)\}$ and $\{p_i^{Wk}(1), P_i^k(1)\}$:

$$G_i : g(\{p_i^{Wk}(0)/G_i\}, P_i^k(0)G_i, q_i, v_i) = g(\{p_i^{Wk}(1)/\Pi_i^k(1)\}, \{P_i^k(1)\}, q_i, v_i), \tag{19}$$

When changes are small, (9) is correct regardless of separability conditions. For discrete changes, the Tornqvist approximation to a Divisia index is justified by (19).

Appendix B: Data Appendix

The study covers 76 trading partners.[54] Bilateral trade flows, measured in thousands of current US dollars, are from CEPII's *Trade, Production and Bilateral Protection Database*[55] (TradeProd) and the United Nation Statistical Division's COMTRADE Database.[56] The TradeProd database is the primary source. The reason is that TradeProd is based on CEPII's *Base pour l'Analyse du Commerce International* (BACI), which implements a consistent procedure for mapping the CIF (cost, insurance and freight) values reported by the importers in COMTRADE to the FOB (free on board) values reported by the exporters in COMTRADE.[57] To further increase the number of non-missing bilateral trade values, we add the

[54] Argentina, Armenia, Australia, Austria, Azerbaijan, Bulgaria, Belgium-Luxembourg, Bolivia, Brazil, Canada, Chile, China, Colombia, Costa Rica, Cyprus, Czech Republic, Germany, Denmark, Ecuador, Egypt, Spain, Estonia, Finland, France, United Kingdom, Greece, Guatemala, Hong Kong, China, Hungary, Indonesia, India, Ireland, Iran, Italy, Jordan, Japan, Kazakhstan, Kenya, Kyrgyz Republic, Korea, Kuwait, Sri Lanka , Lithuania , Latvia, Morocco, Moldova, Mexico, Macedonia, Malta , Mongolia, Mozambique, Mauritius , Malaysia , Netherlands, Norway, Oman, Panama, Philippines, Poland, Portugal, Romania, Russian Federation, Senegal, Singapore , El Salvador, Slovak Republic, Slovenia, Sweden, Trinidad and Tobago, Turkey, Tanzania, Ukraine, Uruguay, United States, Venezuela, South Africa.

[55] For details regarding this database see Mayer, Paillacar and Zignago (2008).

[56] We access COMTRADE through the World Integrated Trade Solution (WITS) software, http://wits.worldbank.org/witsweb/.

[57] As noted in Anderson and Yotov (2010), in principle, gravity theory calls for valuation of exports at delivered prices. In practice, valuation of exports FOB avoids measurement error arising from poor quality transport cost data. For details regarding BACI see Gaulier and Zignago (2008).

mean of the bilateral trade flows from COMTRADE.[58]

Industrial production data comes from two sources. The primary source is the United Nations' UNIDO Industrial Statistics database, which reports industry level output data at the 3 and 4-digit level of ISIC Code (Revisions 2 and 3). In addition to UNIDO, we use CEPII's TradeProd database,[59] as a secondary source.[60] 10.8 percent of the original data were missing after combining the two data sets. As output data are crucial for the calculation of the multilateral resistance indexes, we construct the missing values. First, we interpolate the data to decrease the missing values to 8.6 percent.[61] Then, we extrapolate the rest of the missing values using GDP deflator data, which comes from the World Bank's World Development Indicators (WDI) Database.[62]

We generate internal trade and also expenditure data by combining total shipments data and export data:

$$X_{ii}^k = Y_i^k - \sum_{j \neq i} X_{ij}^k. \tag{20}$$

For expenditures

$$E_j^k = \sum_{i \neq j} (X_{ij}^k + X_{jj}^k). \tag{21}$$

This procedure can result in negative expenditure and internal trade values, and does so for 1.7% of the internal trade observations and for 0.29 percent of the expenditures. In addition, 4.6 percent of the expenditures were missing.[63] To construct the missing expenditure values, first, we interpolate the data, then, we extrapolate the rest using CPI data from the WDI Database.[64]

[58]We also experiment by just using the export data from COMTRADE and then assigning missing trade values to the observations when only data on imports are available. Estimation results are very similar.

[59]TradeProd uses the OECD STAN Industrial Database as well as UNIDO's IndStat Database.

[60]We experiment with two output variables, based on the main data source, to obtain identical results.

[61]Most of the missing observations are for the early years in the sample (1990-1993) and for the former Soviet republics (e.g. Armenia, Estonia, Lithuania, etc.), which declared independence during the early 90s.

[62]GDP deflator data were not available for Belgium-Luxembourg (BLX). We use Belgium's GDP deflator data to proxy for BLX.

[63]Once again, most of the missing observations are for the early years in the sample (1990-1993) and for the former Soviet republics, which declared independence during the early 90s.

[64]CPI data were not available for Belgium-Luxembourg (BLX). We used Belgium's CPI to proxy for BLX.

To handle the negative internal trade and expenditure values, we use the average internal trade to expenditure ratio for each country across all products to fill in the missing values. This has to be done so that the expenditure shares and shipment shares remain consistent by modifying their values in turn. Specifically, let $K(i)$ denote the set of goods for which, for any country i, $X_{ii}^k > 0$. Aggregate across $k \in K(i)$ to form the aggregate version of (20):

$$Y_i^{'} - \sum_{k \in K(i), j \neq i} X_{ij}^k = X_{ii}^{'}. \tag{22}$$

Similarly form 'aggregate' expenditure

$$E_i^{'} = \sum_{k \in K(i), j \neq i} (X_{ji}^k + X_{ii}^k). \tag{23}$$

From these restricted aggregates, form the average ratio of internal trade to expenditure:

$$s_{ii} = X_{ii}^{'}/E_i^{'}, \forall i. \tag{24}$$

Finally, generate the value of inferred internal trade as

$$X_{ii}^k = s_{ii} E_i^{'}, \forall k \notin K(i). \tag{25}$$

Using the generated values from (25), replace the values of internal trade where (20) gives a non-positive value. Then use (20) again with the new data. For consistency of the data, this means that the original data on Y_i^k must be increased by the inferred value of internal trade from (25).

In order to calculate the multilateral resistance indexes, we need data on elasticities of substitution at the commodity level. These data are obtained from Broda et al (2006), who estimate and report 3-digit HS indexes for 73 countries for the period 1994-2003. This period almost coincides with the period of investigation in our study. We use imports as weights to aggregate the original indexes to the level of commodity aggregation in our study. In addition, as some of the original numbers seem implausible,[65] we bound the originals in the interval [4,12] before aggregation. We view the aggregated elasticity indexes as plausible. More homogeneous categories (e.g. Wood and Paper products) have higher values, while less homogeneous categories (e.g. Furniture and Transportation products) have lower numbers.

We experiment with several distance variables based on different approaches in the calculation of internal as well as bilateral distances.[66] Provided that our results are robust to the choice of alternative distance measures, we employ the procedure for distance calculation from Mayer and Zignago (2006).[67] Their approach is appealing because it can be used to calculate consistently both internal distances and bilateral distances.[68] In addition, we follow Eaton and Kortum (2002) to decompose distance effects into four intervals. The distance intervals, in kilometers, are: [0, 3000); [3000, 7000); [7000, 10000); [10000, maximum]. Data on other standard gravity variables such as common language, common border, and colonial ties are from CEPII's *Distances* Database and from Rose (2004).[69] We also generate a set of border dummy variables, which take a value of one for internal trade.

[65] For example, the elasticity estimate for the 3-digit HS commodity category 680, which includes Articles of asphalt, Panels, Boards, Tiles, Blocks, Friction materials etc. is 195.95, while the estimate for category 853 including Electrical capacitors, Electrical resistors, Electric sound/visual signalling equipment etc. is 1.07.

[66] Head and Mayer (2000) provide a nice summary and discussion of alternative distance calculations.

[67] Their procedure is based on Head and Mayer (2000), using the following formula to generate weighted distances: $d_{ij} = \sum_{k \in i} \frac{pop_k}{pop_i} \sum_{l \in j} \frac{pop_l}{pop_j} d_{kl}$, where pop_k is the population of agglomeration k in trading partner i, and pop_l is the population of agglomeration l in trading partner j, and d_{kl} is the distance between agglomeration k and agglomeration l, measured in kilometers, and calculated by the Great Circle Distance Formula. All data on latitude, longitude, and population is from the World Gazetteer web site.

[68] In the few instances where we were not able to implement Mayer and Zignago's procedure, we just took the distance between the main cities from the two trading partners.

[69] Rose's (2004) original data covers the period up to the year of 2000, so we update some of the variables in order to match the time span investigated in this study.