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Robert C. Feenstra
John Romalis

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ABSTRACT

The unit values of internationally traded goods are heavily influenced by quality. We model this in an extended monopolistic competition framework where, in addition to choosing price, firms simultaneously choose quality. We allow countries to have non-homothetic demand for quality. The optimal choice of quality by firms reflects this non-homothetic demand as well as the costs of production, including specific transport costs, under the "Washington apples" effect. We estimate the implied gravity equation using detailed bilateral trade data for about 200 countries over 1984-2008. Our system identifies quality and quality-adjusted prices, from which we will construct price indexes for imports and exports for each country that will be incorporated into the next generation of the Penn World Table.

Robert C. Feenstra
Department of Economics
University of California, Davis
One Shields Avenue
Davis, CA 95616
and NBER
rcfeenstra@ucdavis.edu

John Romalis
Australian National University
Research School of Economics
College of Business and Economics
Crisp Building 26, Room 2078
The Australian National University ACT 0200
Australia
and NBER
jromalis@gmail.com

1. Introduction

It has long been known that the unit values of internationally traded goods are heavily influenced by their quality (Kravis and Lipsey, 1974). That is the reason why import and export price indexes for the United States and many other countries no longer use any unit-value information, but instead rely on price surveys from trading firms. Likewise, when making international comparisons of real GDP, researchers such as Summers and Heston (1991) rely on the prices surveys of the International Comparisons Program, which collects prices of identical products across countries. Those prices are only collected for final goods sold in each country, however, and are then used to construct real GDP in the Penn World Table (PWT). Recently, it has been proposed that PWT could be extended to incorporate the prices of exports and imports, which would allow a distinction to be made between real GDP from the consumers and producers points of view: these differ by the terms of trade faced by countries (Feenstra et al, 2009). In order to make this distinction we need to have quality-adjusted prices (or unit values) for a wide range of traded goods over many countries and years. That is the goal of our study.

To achieve this goal, we extend the model of Melitz (2003) to allow for endogenous quality choice by firms.[1] We are not the first to attempt to disentangle quality from trade unit values, and other recent authors to do so include Schott (2004, 2008), Hallak (2006), Hallak and Schott (2011), Khandelwal (2010) and Martin and Méjean (2010).[2] These studies rely on the demand side to identify quality. In the words of Khandelwal (2010, p. 1451): "The procedure

[1] Baldwin and Harrigan (2011) argue that introducing quality in the Melitz model is essential to make it consistent with empirical observations, but in their framework quality is exogenous. Models with endogenous quality choice by heterogeneous firms include Gervias (2010), Khandelwal (2010) and Mandel (2009). The latter two paper have simultaneous choice of price and quality, as we use here. In contrast, Gervias has quality chosen for the lifetime of a product. This yields a solution where quality is proportional to firm productivity, thereby providing a micro-foundation for this assumption made in Baldwin and Harrigan (2011).

[2] Another line of literature empirically distinguishes between productivity and quality versions of the Melitz (2003) model: see Baldwin and Ito (2011), Crozet, Head and Mayer (2012), Johnson (2012) and Mandel (2009).

utilizes both unit value and quantity information to infer quality and has a straightforward intuition: conditional on price, imports with higher market shares are assigned higher quality." Likewise, Hallak and Schott (2011) rely on trade balances to identify quality. To this demand-side information we will add a supply side, drawing on the well-known "Washington apples" effect (Alchian and Allen, 1964; Hummels and Skiba, 2004): goods of higher quality are shipped longer distances. We will find that this positive relationship between exporter f.o.b. prices and distance is an immediate implication of the first-order condition of firms for optimal quality choice. This first-order condition gives us powerful additional information from which to identify quality.

In section 2 we specify our model, where firms in each country *simultaneously* choose price and quality. That is, we are thinking of quality characteristics as being modified easily and tailored to each market: the specification of a Volkswagen Golf sold in various countries is a realistic example. Like Hallak (2006) and Fajgelbaum, Grossman and Helpman (2011a,b), countries have non-homothetic demand for quality. As in Verhoogen (2008), we assume a Cobb-Douglas production function for quality where firms differ in their productivities. Then solving the firm's problem, we find that quality is a simple log-linear function of productivity and wages, as well as the specific transport costs to the destination market and that country's valuation of quality. Specializing to the CES demand system, we solve for the prices charged by firms and find that an exporter's f.o.b. prices are directly proportional to specific transport costs, illustrating the Washington apples effect. It follows that log quality is proportional to the log of the exporter's f.o.b. price divided by the productivity-adjusted wages.

In order to implement this measure of quality, we therefore need accurate information on wages and other input prices as well as the productivity of exporters to each destination market. Verhoogen (2008) argues that multiple factors are needed to produce high-quality outputs, and De Loeker and Warzynski (2011) likewise argue that it is important to model all the inputs used by a firm in order to measure productivity, especially for exporters. The ability to obtain data on such input prices for a broad range of industries (i.e. every merchandise export) and countries (i.e. all countries included in the Penn World Table) is a formidable challenge. To overcome this challenge, we rely on the equilibrium assumption that the marginal exporting firm to each destination market earns zero profits, as in Melitz (2003). We further assume that the distribution of productivities across firms is Pareto. Then we can use the zero-cutoff-profit condition to solve for the productivity-adjusted wages and firm-level quality.

Because our goal is to estimate quality for many goods and countries we do not rely on firm-level data, but in section 3 aggregate to the disaggregate industry level, in which case the c.i.f. and f.o.b. prices are measured by unit values. We derive the gravity equation and show that it includes new terms as compared to Chaney (2008), due to endogenous quality. In sections 4 and 5, we estimate the gravity equation using detailed bilateral trade data at the 4-digit SITC digit level (nearly 800 products per year) for about 200 countries during 1984-2008. Our median estimate of the elasticity of substitution is higher than that in Broda and Weinstein (2006) due to our expanded sample over many countries along with the fact that quality is included, and by using a specification that is more robust to measurement error. Our median estimate of the Pareto parameter in quite close to Eaton and Kortum (2002), who also consider trade between many countries.

Given the parameter estimates*,* product quality is readily constructed in section 6. Our results broadly conform to our expectations. Developed countries export and import higher quality goods than do poorer countries, confirming results of Schott (2004) and Hallak (2006), but the quality-adjusted prices vary much less than the raw unit values. Countries' quality-adjusted terms of trade are negatively related to their level of development. We provide indexes of quality and quality-adjusted prices at the 4-digit SITC and 1-digit Broad Economic Categories (distinguishing food and beverages, other consumer goods, capital, fuels, intermediate inputs and transport equipment), that should be useful to researchers interested in the time-series or cross-country properties of these indexes and that will be incorporated into the next generation of the PWT (Feenstra, Inklaar and Timmer, 2012).

2. Optimal Quality Choice

Consumer Problem

Suppose that consumers in country k have available i=1,…,N varieties of a differentiated product in a sector. These products can come from different source countries. We should really think of each variety as indexed by the triple *(i,j,t)*, where i is the country of origin, j is the firm and t is time. But initially, we will simply use the notation i for product varieties. We denote the price and quality of good i in country k by p_i^k and z_i^k, respectively. We suppose that the demand in country k arises from the expenditure function $E^k = E(p_1^k / z_1^{\alpha^k}, ..., p_N^k / z_N^{\alpha^k}, U^k),$ where quality z_i^k is raised to the power $\alpha^k > 0$, which we denote by $z_i^{\alpha^k} \equiv (z_i^k)^{\alpha^k}$ for brevity. Thus, quality acts as a shift parameter in the expenditure function. Hallak (2006) has introduced similar exponents on quality, but in the context of the *direct* utility function. In that case it is not possible to makes the exponents α^k depend on utility or per-capita income; but by working with the expenditure function we will be able to do just that.

Differentiating the expenditure function to compute demand q_i^k:

$$q_i^k = \frac{\partial E}{\partial p_i^k} = \frac{E_i}{z_i^{\alpha^k}},$$

where E_i denotes the derivative of E with respect to its i$^{\text{th}}$ argument. Denoting this i$^{\text{th}}$ argument by the *quality-adjusted prices* $P_i^k \equiv p_i^k / z_i^{\alpha^k}$, and defining $Q_i^k \equiv z_i^{\alpha^k} q_i^k$ as the *quality-adjusted demand*, we can re-arrange terms above to obtain $Q_i^k = E_i(P_1^k, ..., P_N^k, U^k)$, so that working with the quality-adjusted magnitudes still gives quantity as the derivative of the expenditure function with respect to price.

We can generalize the expenditure function by allowing the exponents $\alpha^k = h(U^k)$ to depend on utility, so that:

$$E^k = E[p_1^k / z_1^{h(U^k)}, ..., p_N^k / z_N^{h(U^k)}, U^k], \tag{1}$$

where $z_i^{h(U^k)} \equiv (z_i^k)^{h(U^k)}$. Because $\alpha^k = h(U^k)$ depends on utility, this expenditure function has non-homothetic demand for quality, as in Fajgelbaum, Grossman and Helpman (2011a,b).[3] This is a valid expenditure function provided that it is increasing in utility and non-decreasing in price, which can be readily checked for a specific functional form.[4] For most of the paper we shall rely on a "non-homothetic CES" expenditure function, defined over a continuum of products i as,

$$E^k = U^k \left[\int_i \left(p_i^k / z_i^{h(U^k)} \right)^{(1-\sigma)} di \right]^{\frac{1}{(1-\sigma)}}, \tag{2a}$$

with, $$h(U^k) = 1 + \lambda \ln U^k, \quad \text{for } U^k > 0. \tag{2b}$$

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[3] Other recent literature including Bekkers *et al* (2010), Choi *et al* (2009) and Simonovska (2011) analyze models of international trade and quality where non-homothetic demand plays a central role.

[4] The idea of allowing the parameters of the expenditure function to depend on utility is borrowed from Deaton and Muellbauer (1980, pp. 154-158), who define an expenditure function as a utility-weighted combination of any two functions that are non-decreasing in price, which is valid provided that the resulting function is increasing in utility. Deaton and Muellbauer use this approach to obtain a class of expenditure functions where aggregate demand depends in a simple way on the moments of the income distribution.

Totally differentiating (2a) with respect to utility using $E_i = Q_i^k$ and (2b), we obtain:

$$\frac{\partial E^k}{\partial U^k} = \frac{E^k}{U^k} + \int_i Q_i^k \frac{dP_i^k}{dU^k} di = \frac{E^k}{U^k}\left[1 - \lambda \int_i \left(\frac{P_i^k Q_i^k}{E^k}\right) \ln z_i^k \, di\right],$$

since $dP_i^k / dU^k = -P_i^k \ln z_i^k h'(U^k)$ and $\lambda = U^k h'(U^k)$. The final integral above is interpreted as the average of log quality across products. In our empirical work we will model $\alpha^k = h(U^k)$ as depending on per-capita income rather than unobserved utility. We see that the non-homothetic CES expenditure function will be increasing in utility provided that λ is sufficiently small, which is readily confirmed in our estimates.

Firms' Problem

We now add the subscript *j* for firms, while *i* denotes their country of origin and *k* the destination, so that *(i,j,k)* denotes a unique variety. Firms make the optimal choice of the quality z_{ij}^k to send to country *k*. We suppose there are both specific and iceberg trade costs between the countries. The specific trade costs is given by T_i^k, which can depend on factor prices and the distance to the destination market *k*.[5] The iceberg trade cost is denoted by τ_i^k, so that τ_i^k units of the good are exported in order for one unit to arrive. We assume that the iceberg costs, which includes one plus the *ad valorem* tariff denoted by tar_i^k, are applied to the value *inclusive* of the specific trade costs.[6] Then letting p_i^{*k} denote the exporters' f.o.b. price, the tariff-inclusive c.i.f. price is $p_i^k \equiv \tau_i^k (p_i^{*k} + T_i^k)$, and the net-of-tariff c.i.f. price is p_i^k / tar_i^k.

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[5] That is, we could write $T_i^k = w_i d_i^k$, where d_i^k is in units of the aggregate factor and depends on distance.

[6] Most countries apply tariffs to the transport-inclusive (c.i.f.) price of a product. The exceptions are Afghanistan, Australia, Botswana, Canada, Democratic Republic of the Congo, Lesotho, Namibia, New Zealand, Puerto Rico, South Africa, Swaziland, and the United States. See the Customs Info Database at http://export.customsinfo.com/ and http://export.gov/logistics/eg_main_018142.asp.

We assume that output is produced with a composite input whose quantity is denoted by L_{ij}^k. The amount of the composite input needed for one unit of production is denoted l_{ij}^k, and the total input requirements (or inverse of the production function) takes the form:

$$L_{ij}^k = l_{ij}^k y_{ij}^k + f_{ij}^k ,$$

where y_{ij}^k is the output sold by firm *j* to country *k*, and f_{ij}^k are the fixed costs of selling there. We will not specify these fixed costs until the next section, focusing here on the variable input. In order to produce one unit of a good with product quality z_{ij}^k the firm must use the variable input l_{ij}^k, according to the Cobb-Douglas production function:

$$z_{ij}^k = (l_{ij}^k \varphi_{ij})^{\theta} , \qquad (3)$$

where $0 < \theta < 1$ reflects diminishing returns to quality and φ_{ij} denotes the productivity of firm *j* in country *i*. We think of l_{ij}^k as an aggregate of labor inputs, such as high and low-skilled labor and entrepreneurial ability, as in Verhoogen (2008), and denote its price by the wage w_i. The marginal cost of producing a good of quality z_{ij}^k is then solved from (3) as,

$$c_{ij}(z_{ij}^k, w_i) = w_{it} x_{ij}^k = w_i (z_{ij}^k)^{1/\theta} / \varphi_{ij} . \qquad (4)$$

Firms simultaneously choose f.o.b. prices p_{ij}^{*k} and characteristics z_{ij}^k for each destination market. From the iceberg costs, τ_i^k units of the good are exported in order for one unit to arrive, so total exports are $y_{ij}^k = \tau_i^k q_{ij}^k$. When evaluating profits from exporting to country *k,* we need to divide by one plus the *ad valorem* tariff tar_i^k, obtaining:

$$\begin{aligned} \max_{p_{ij}^{*k}, z_{ij}^k} \; [p_{ij}^{*k} - c_{ij}(z_{ij}^k, w_i)] \frac{\tau_i^k q_{ij}^k}{tar_i^k} &= \max_{p_{ij}^{*k}, z_{ij}^k} \left[\frac{p_{ij}^{*k}}{z_{ij}^{\alpha^k}} - \frac{c_{ij}(z_{ij}^k, w_i)}{z_{ij}^{\alpha^k}} \right] \frac{\tau_i^k Q_{ij}^k}{tar_i^k} \\ &= \max_{P_{ij}^k, z_{ij}^k} \left\{ P_{ij}^k - \tau_i^k \frac{[c_{ij}(z_{ij}^k, w_i) + T_i^k]}{z_{ij}^{\alpha^k}} \right\} \frac{Q_{ij}^k}{tar_i^k} . \end{aligned} \qquad (5)$$

The first equality in (5) converts from observed to quality-adjusted consumption, while the second line converts to quality-adjusted, tariff-inclusive c.i.f. prices $P_{ij}^{k} \equiv \tau_i^k (p_{ij}^{*k} + T_i^k) / z_{ij}^{\alpha^k}$, while changing the choice variables from p_{ij}^{*k}, z_{ij}^{k} to P_{ij}^{k}, z_{ij}^{k}. This change in variables relies on the form of the expenditure function in (1) and our assumption that *prices and characteristics are chosen simultaneously,* since we are thinking of quality characteristics that can be changed easily, but does not rely on the CES functional form in (2).

It is immediate that to maximize profits in (5), the firms must choose z_{ij}^{k} to minimize $[c_{ij}(z_{ij}^{k}, w_i) + T_i^k] / z_{ik}^{\alpha^k}$. In the case where $\alpha^k = 1$, this problem is interpreted as *minimizing the average variable cost per unit of quality, inclusive of specific trade costs,* which is obtained where marginal cost equals average cost as found by Rodriguez (1979). More generally, with $\alpha^k > 0$ the solution to this problem is:

$$\frac{\partial c_{ij}(z_{ij}^{k}, w_i)}{\partial z_{ij}^{k}} = \alpha^k \frac{[c_{ij}(z_{ij}^{k}, w_i) + T_i^k]}{z_{ij}^{k}}, \tag{6}$$

so there is a wedge of α^k between the marginal and average costs of producing quality. The second-order condition for this minimization problem is satisfied if and only if $\partial^2 c_{ij} / \partial (z_{ij}^{k})^2 > 0$, so there must be increasing marginal costs of improving quality. In that case, either an increase in the valuation of quality α^k or an increase in the specific transport costs to the destination market T_i^k will *raise* quality z_{ij}^{k}. This occurs in particular with an increase in T_i^k due to greater distance, which is the well-known "Washington apples" effect.

Making use of the Cobb-Douglas production function for quality in (3), and associated cost function in (4), the second-order conditions are satisfied when $0 < \theta < 1$ which we have already assumed. The first-order condition (6) can be solved for quality:

$$\ln z_i^k = \theta\left[\ln T_i^k - \ln(w_i / \varphi_{ij}) + \ln(\alpha^k\theta / (1-\alpha^k\theta))\right], \tag{7}$$

where we assume that $\alpha^k\theta < 1$. Conveniently, the Cobb-Douglas production function and specific trade costs give us a log-linear form for the optimal quality choice. Since we are allowing $\alpha^k = h(U^k)$ to depend on the utility of the destination market, it follows that richer countries (with higher utility) will import higher quality, as found empirically by Hallak (2006). In addition, quality in (7) is rising in the productivity of the exporting firm, confirming the finding of Schott (2004) that richer (more productive) countries export higher quality goods.[7] Substituting (7) into the cost function (4), we obtain $c_{ij}(z_{ij}^k, w_i) = [\alpha^k\theta / (1-\alpha^k\theta)]T_i^k$. Thus, the marginal costs of production are proportional to the specific trade costs, as we shall use below.

Now suppose that the CES expenditure function in (2) applies. Solving (5) for the optimal choice of the f.o.b. price, we obtain the familiar markup,

$$(p_{ij}^{*k} + T_i^k) = [c_{ij}(z_{ij}^k, w_i) + T_i^k]\left(\frac{\sigma}{\sigma-1}\right).$$

This equation shows that firms not only markup over marginal costs c_{ij} in the usual manner, they also markup over specific trade costs. Then using the relation $c_{ij}(z_{ij}^k, w_i) = [\alpha^k\theta / (1-\alpha^k\theta)]T_i^k$, we readily solve for the f.o.b. and tariff-inclusive c.i.f. prices as:

$$p_{ij}^{*k} = T_i^k\left[\left(\frac{1}{1-\alpha^k\theta}\right)\left(\frac{\sigma}{\sigma-1}\right)-1\right] \equiv \overline{p_i^{*k}}, \tag{8a}$$

$$p_{ij}^{k} = \tau_i^k T_i^k\left[\left(\frac{1}{1-\alpha^k\theta}\right)\left(\frac{\sigma}{\sigma-1}\right)\right] \equiv \overline{p_i^{k}}. \tag{8b}$$

Thus, both the f.o.b. and c.i.f. prices vary across destination markets *k* in proportion to the

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[7] As mentioned in note 5, we could write $T_i^k = w_i d_i^k$, where d_i^k is in units of the aggregate factor and depends on distance. In that case, wages w_i (which also depend on productivity) cancel out from (7).

specific transport costs to each market, but are independent of the productivity of the firm *j*, as indicated by the notation $\overline{p_i^{*k}}$ and $\overline{p_i^k}$. This result is obtained because more efficient firms sell higher quality goods, leading to constant prices in each destination market.[8]

Both the f.o.b. and c.i.f. prices are increasing in the destination country's preference for quality, α^k. This provides us with a method to estimate these preferences using data on f.o.b. unit values $\ln uv_i^{*k} = \ln \overline{p_i^{*k}} + u_i^{*k}$, with measurement error u_i^{*k}. We model α^k as depending on real GDP per capita of country *k* from the Penn World Table. Taking logs of (8a), adding a time subscript *t* and a SITC goods subscript *g,* and assuming that specific transport costs depend on distance, we estimate:

$$\ln uv_{igt}^{*k} = \delta_{igt} + \beta_g \ln(\text{dist}_i^k) + \ln\left[\left(\frac{1}{1-\alpha_{gt}^k \theta_g}\right)\left(\frac{\sigma_g}{\sigma_g - 1}\right) - 1\right] + u_{igt}^{*k}, \qquad (9a)$$

with,
$$\alpha_{gt}^k = 1 + \lambda_g \ln\left(RGDPL_t^k \,/\, RGDPL_t^{US}\right), \qquad (9b)$$

where δ_{igt} is a source country-time fixed effect. We measure real GDP per capita, $RGDPL_t^k$, relative to that in the United States as a normalization. Substituting (9b) into (9a), and using preliminary estimates of σ_g and θ_g that we shall describe later, we can obtain estimates of λ_g using nonlinear least squares. This regression is run for each SITC 4-digit industry over 1984-2008, with the results shown in Figures 1 and 2.

In Figure 1 we shown the frequency of estimates for β_g, the coefficient on log distance. Its median value over 862 4-digit SITC industries is 0.15. Only 2.5% of the estimates are significantly negative. The fact that the f.o.b. unit value – which is *net* of transport costs – is

[8] Our result is a razor-edge case between having the largest firms charge low prices (due to high productivity) or high prices (due to high quality) in a given destination market. Other authors have distinguished those two cases using firm-level data: see note 2. While this razor-edge case simplifies our analytical results, such as taking averages in section 3, it is not essential to our analysis because we ultimately rely on industry rather than firm-level prices.

Figure 1: Frequency Distribution for Estimates of β_g

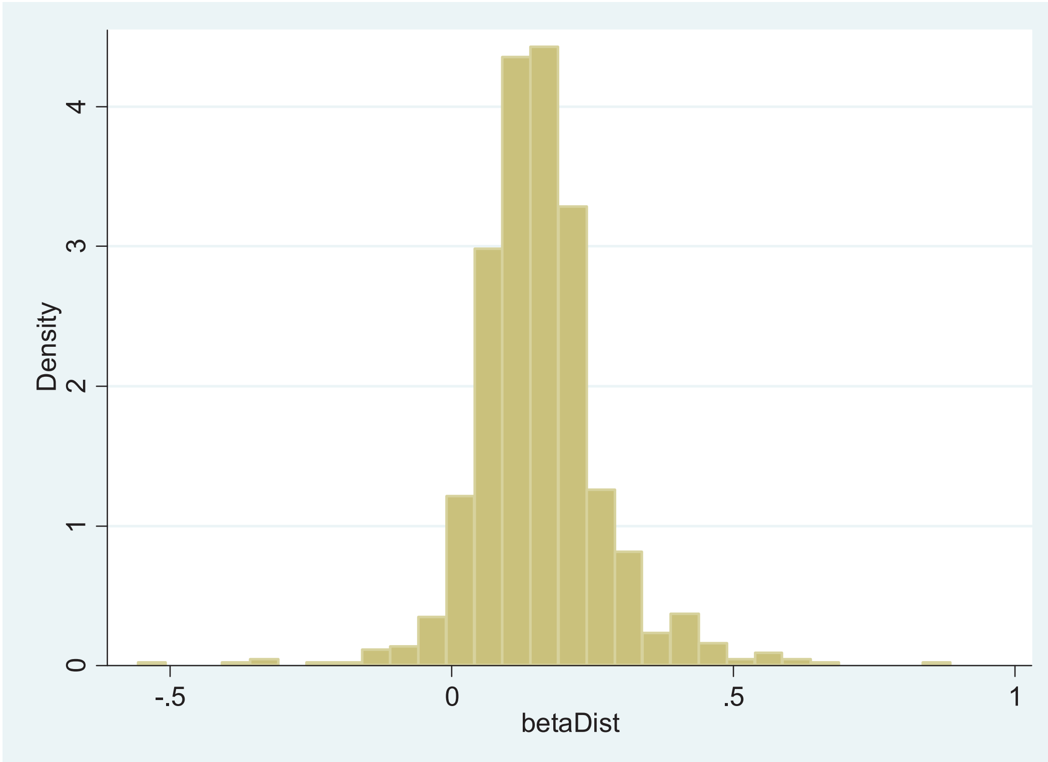
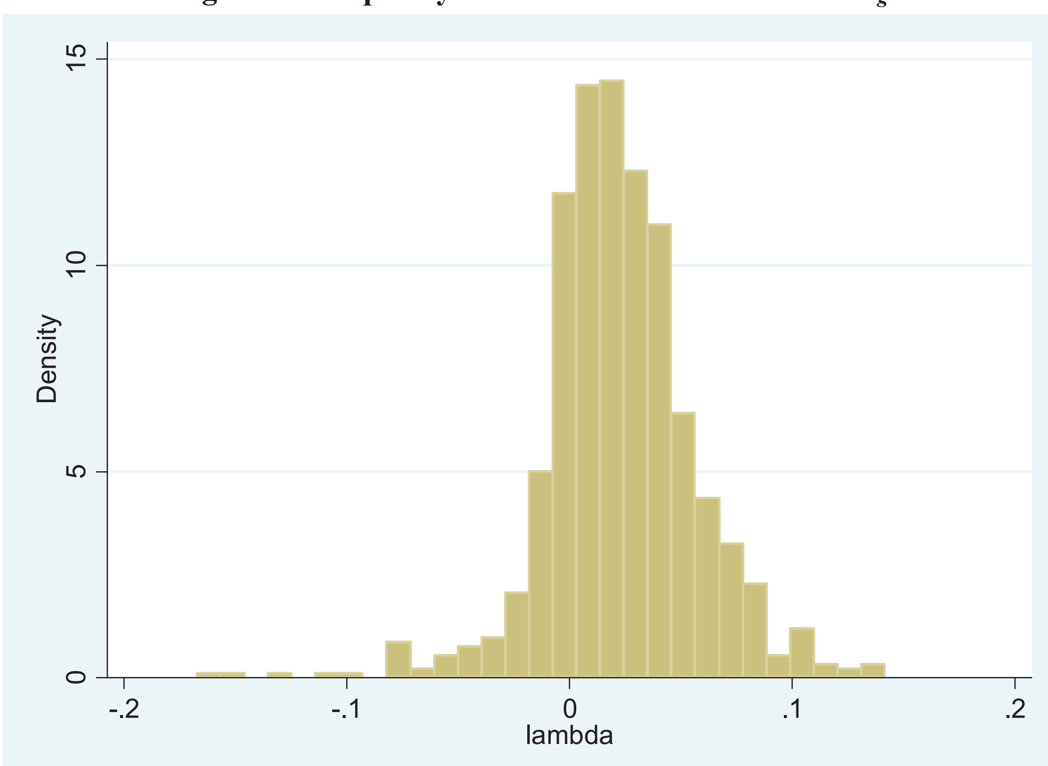


Figure 2: Frequency Distribution for Estimates of λ_g



increasing in distance is interpreted by Hummels and Skiba (2004) as evidence of the "Washington apples" effect, whereby quality grows with distance. In fact, in our model log quality is only a fraction of the log f.o.b. price, as shown by combining (7) and (8a) to obtain (without the *g* and *t* subscripts):

$$\ln z_i^k = \theta\left[\ln(\kappa_1^k \overline{p_i^{*k}}) - \ln(w_i / \varphi_{ij})\right], \text{ with } \kappa_1^k \equiv \left[\frac{\alpha^k \theta(\sigma-1)}{1+\alpha^k \theta(\sigma-1)}\right]. \tag{10}$$

To isolate quality from the f.o.b. price we need to know the key parameter θ from the production function for quality, which we estimate in section 4.

In Figure 2 we show the frequency distribution for estimates of λ_g , the coefficient of real GDP per capita in determining α_{gt}^k. Its median value over the 4-digit SITC industries is 0.021, and about 14% of the estimates are significantly negative. We will replace the negative but insignificant estimates by zero, but do not alter the negative and significant estimates: there are plausible cases where lower-income countries prefer higher quality due to the changing composition of goods within SITC 4-digit categories. The leading example is SITC 3341, "Gasoline and other Light Fuels," which includes fuels for aircraft. It has $\lambda_g = -0.07$, the largest significant negative value, since many small, low-income economies (especially island countries) without refining capacity require relatively more of the higher-quality aircraft fuel. The implied values for α_{gt}^k range between 0.42 and 1.31 over all goods.

For quality in (10), the corresponding quality-adjusted price $P_{ij}^k = \overline{p_i^k} / z_{ij}^{\alpha^k}$ is:

$$P_{ij}^k = \overline{p_i^k}\left[(w_i / \varphi_{ij}) \Big/ \kappa_1^k \overline{p_i^{*k}}\right]^{\alpha^k \theta}. \tag{11}$$

Since from (8) the c.i.f. and f.o.b. prices do not differ across firms selling to each destination market, it follows that the quality-adjusted price is decreasing in the productivity φ_{ij} of the exporter, but at a rate that differs from Chaney (2008). For the CES expenditure function in (2), sales depend on the quality-adjusted price with elasticity $(1-\sigma)$, and from (11), the price depends on productivity with elasticity $-\alpha^k\theta$, so that firms' sales depends on productivity with elasticity $\alpha^k\theta(\sigma-1)$. Below we will assume a Pareto distribution for productivities with parameter γ. It follows that firms' sales in our endogenous-quality model are Pareto distributed with parameter $\zeta^k = \gamma / [\alpha^k\theta(\sigma-1)]$ in country *k*. We can see that in order for our model to mimic the Melitz-Chaney model, we need to have $\alpha^k\theta \to 1,$ so that prices would decline at the rate $\gamma / (\sigma-1)$ as in Chaney (2008). Setting $\alpha^k\theta = 1$ is not permitted because the quality approaches infinity in (7), but we will occasionally let $\alpha^k\theta \to 1$ to compare our results to the Melitz-Chaney model.

3. Solving for Productivity-Adjusted Wages

As discussed in section 1, it would be a formidable challenge to assemble the data on wages, other input prices and firms' productivities needed to directly measure quality in (10) across many goods and countries. In our trade data we will not have firm-level information. Accordingly, we rely instead on the zero-cutoff-profit condition of Melitz (2003) to solve for the productivity-adjusted wage of the marginal exporter to each destination market. In addition, we shall aggregate prices and quality to the industry level to obtain observable magnitudes, which will turn out to be useful in solving for the marginal exporter.

Zero-Cutoff-Profit Condition

We let $\hat{\varphi}_i^k$ denote the cutoff productivity for a firm in country *i* that can just cover the fixed costs of exporting to country *k*. Using this productivity in (11), $\hat{P}_i^k$ denotes the quality-adjusted price for the marginal exporter:

$$\hat{P}_i^k = \overline{p_i^k}\left[(w_i / \hat{\varphi}_i^k) \big/ \kappa_1^k \overline{p_i^{*k}}\right]^{\alpha^k \theta} . \tag{11'}$$

We let $\hat{Q}_i^k$ denote the quantity of exports for this marginal firm so that $\hat{X}_i^k \equiv \hat{P}_i^k \hat{Q}_i^k$ is tariff-inclusive export revenue. From the CES markups, profits earned by the firm in (5) are then $(\hat{X}_i^k / tar_i^k \sigma)$, which must cover fixed costs in the zero-cutoff-profit (ZCP) condition:

$$\frac{\hat{X}_i^k}{tar_i^k \sigma} = \left(\frac{w_i}{\hat{\varphi}_i^k}\right) f_i^k . \tag{12}$$

There are two features of this ZCP condition that deserve attention. First, one plus the *ad valorem* tariff tar_i^k appears in the denominator on the left because tariffs must be deducted from revenue before computing profits. Equivalently, we can move the term tar_i^k to the right where it will multiply fixed costs f_i^k, which is how that term will appear in the ensuing formulas.

Second, we have written wages on the right of (12) as adjusted for productivity of the ZCP exporter. That is, we are assuming that an exporting firm's productivity applies equally well to variable and fixed costs. We make this assumption because export revenue depends on the quality-adjusted price $\hat{P}_i^k$ (in addition to the CES price index, specified below), so the solution for the quality-adjusted price from (12) is very sensitive to the specification of fixed costs. By using the productivity-adjusted wages in (12), the solution for the quality-adjusted prices depends on *more than* just the fixed cost f_i^k. This assumption is also made by Bilbiie, Ghironi, and Melitz (2012), though in their case, productivity is equal across firms.

To make use of (12) we need to aggregate across all firms with higher productivity than the marginal exporter, obtaining total sectoral exports from country *i* to *k.* In addition, following Melitz (2003) we form the CES averages of the quality-adjusted prices in (11). To perform these aggregations, we add the assumption that productivity is Pareto distributed with cumulative distribution $G_i(\varphi) = 1-(\varphi / \varphi_i)^{-\gamma}$, where the location parameter $\varphi_i \leq \varphi$ denotes the lower-bound to the productivities of firms in country *i.* By varying this lower-bound we can achieve differences in average productivity across countries, which is realistic, but for analytical convenience we assume that the dispersion parameter γ is identical across countries.[9] The density function is $g_i(\varphi) = \gamma\varphi^{-(\gamma+1)}\varphi_i^{\gamma}$, and the density *conditional* on exporting to country *k* is $g_i(\varphi)/[1-G_i(\hat{\varphi}_i^k)]$, for $\varphi \geq \hat{\varphi}_i^k$.

In order to aggregate over exporters, we note that the ratio of demand for firm *j* and the cutoff firm, both exporting to the same destination market *k*, is $Q_{ij}^k / \hat{Q}_i^k = (P_{ij}^k / \hat{P}_i^k)^{-\sigma}$, so that relative firm revenue is $X_{ij}^k / \hat{X}_i^k = (P_{ij}^k / \hat{P}_i^k)^{1-\sigma}$. Denoting the mass of firms in country *i* by M_i, total exports in this sector from country *i* to *k* are:

$$\begin{aligned} X_i^k &\equiv M_i \int_{\hat{\varphi}_i^k}^{\infty} X_{ij}^k g_i(\varphi)d\varphi = M_i \int_{\hat{\varphi}_i^k}^{\infty} \hat{X}_i^k (P_{ij}^k / \hat{P}_i^k)^{1-\sigma} g_i(\varphi)d\varphi \\ &= M_i \hat{X}_i^k \left[\int_{\hat{\varphi}_i^k}^{\infty} \left(\frac{\hat{\varphi}_i^k}{\varphi} \right)^{\alpha^k \theta(1-\sigma)} g_i(\varphi)d\varphi \right] = \hat{X}_i^k M_i \left(\frac{\hat{\varphi}_i^k}{\varphi_i} \right)^{-\gamma} \left[\frac{\gamma}{\gamma - \alpha^k \theta(\sigma-1)} \right], \end{aligned} \tag{13}$$

as obtained by evaluating the integral and assuming $\gamma > \alpha^k \theta(\sigma-1)$. Substituting from (12) for $\hat{X}_i^k$, we solve for the wage relative to the cutoff productivity:

$$\left(\frac{w_i}{\hat{\varphi}_i^k} \right)^{1+\gamma} = \frac{X_i^k / \kappa_2^k tar_i^k f_i^k}{M_i (\varphi_i / w_i)^{\gamma}}, \text{ with } \kappa_2^k = \left[\frac{\sigma\gamma}{\gamma - \alpha^k \theta(\sigma-1)} \right]. \tag{14}$$

Substituting this solution for productivity-adjusted wages into (11'), we readily obtain an

[9] In this respect we are making the same assumption as in Eaton and Kortum (2002), who allowed for different location parameters of the Frechét distribution across countries, but with the same dispersion parameter.

expression for the quality-adjusted price for the marginal exporter. Our interest is in the CES average of the quality-adjusted prices, which for firms in country *i* exporting to *k* is:

$$\overline{P_i^k} \equiv \left[\int_{\hat{\varphi}_i^k}^{\infty} P_{ij}^k(\varphi)^{(1-\sigma)} \frac{g_i(\varphi)}{\left[1 - G_i(\hat{\varphi}_i^k)\right]} d\varphi \right]^{\frac{1}{1-\sigma}} = \left[\frac{\gamma}{\gamma - \alpha^k \theta(\sigma - 1)} \right]^{\frac{1}{1-\sigma}} \hat{P}_i^k, \qquad (15)$$

as obtained by substituting for $P_{ij}^k(\varphi_{ij})$ from (11) and computing the integral for $\gamma > \alpha^k \theta(\sigma - 1)$. This expression shows that the *average* quality-adjusted price $\overline{P_i^k}$ is proportional to the *cut-off* price $\hat{P}_i^k$, with the factor of proportionality depending only on model parameters.

Combining (11'), (14) and (15), we therefore obtain,

$$\overline{P_i^k} = \left(\overline{p_i^k} \Big/ \left(\kappa_1^k \overline{p_i^{*k}} \right)^{\alpha^k \theta} \right) \left(\frac{X_i^k / \kappa_2^k tar_i^k f_i^k}{M_i (\varphi_i / w_i)^{\gamma}} \right)^{\frac{\alpha^k \theta}{(1+\gamma)}}. \qquad (16)$$

Notably, an increase in exports to a market, given the mass of firms, *raises* the relative quality-adjusted price. That occurs because an increase in relative exports means that less-efficient firms are exporting to that market, and therefore average quality falls. That relationship sounds contrary to the demand-side intuition discussed in section 1: given nominal prices, higher sales to a market should mean higher quality. In fact, that intuition continues to hold in our model, and we shall use it below in conjunction with (16) to solve for the quality-adjusted prices.

Import Demand

Returning to the zero-cutoff-profit condition in (12), while the firm-level sales $\hat{X}_i^k$ are not observed in our data, they equal CES demand from the expenditure function in (2). That is, $\hat{X}_i^k = (\hat{P}_i^k)^{-(\sigma-1)}[Y^k P^{k(\sigma-1)}]$, where P^k is the exact price index corresponding to the CES expenditure function in (2):

$$P^k \equiv \left[\sum_i M_i \int_{\hat{\varphi}_i^k}^{\infty} P_{ij}^k(\varphi)^{(1-\sigma)} g_i(\varphi) d\varphi \right]^{1/(1-\sigma)} = \left[\sum_i \overline{P_i^k}^{(1-\sigma)} M_i \left(\frac{\hat{\varphi}_i^k}{\varphi_i} \right)^{-\gamma} \right]^{1/(1-\sigma)}, \qquad (17)$$

using (15). Also using aggregate exports along with (14), demand is re-expressed as:

$$\left(\frac{X_i^k / \kappa_{2i}^k f_i^k}{M_i (\varphi_i / w_i)^\gamma} \right) = \left(\frac{\overline{P_i^k}}{P^k} \right)^{-(\sigma-1)(1+\gamma)} \left(\frac{Y^k}{\kappa_2^k tar_i^k f_i^k} \right)^{(1+\gamma)}. \qquad (18)$$

Higher exports on the left of this expression imply a lower quality-adjusted price on the right, *ceteris paribus,* so this equation has the demand-side intuition. The mass of potential exporters M_i enters this equation because if there are more firms selling from country *i* to *k* then exports will be higher. The presence of this term complicates all demand-side attempts to measure quality, because *either* a greater mass of firms (leading to more variety) or higher quality (leading to lower quality-adjusted prices) will raise exports in (18). This problem is dealt with in different ways by Hallak and Schott (2011), Hummels and Klenow (2005) and Khandelwal (2010): the latter author, for example, uses exporting country population to measure the mass of exporters. We will rely on a similar assumption in our estimation, but first show how the mass of exporters can be eliminated by combining the demand-side relation (18) with the supply-side relation (16) to solve for the quality-adjusted price and for exports.

Gravity Equation and Quality-Adjusted Prices

Combining (16) and (18) we readily solve for exports X_i^k, which is the gravity equation. To present this solution in the most compact form, we use the techniques of Chaney (2008) to solve for the CES price index in (17) (see Appendix A for details). We also assume a specific functional form for the fixed costs:

$$f_i^k = \left(\frac{Y^k}{uv^k}\right)^{\beta_0} \exp\left(\sum_{n=1}^{4} \beta_n F_{ni}^k\right). \qquad (19)$$

The first term in (19) follows the hypothesis of Arkolakis (2010) that small markets have lower fixed costs because it easier to reach all customers: uv^k is a deflator for import expenditure in the industry in question constructed from import unit values, so that (Y^k / uv^k) measures quantity. Arkolakis allows the fixed costs f_i^k to depend on $(Y^k / uv^k)^{\beta_0}$, where β_0 is a parameter indicating the sensitivity to market size. This parameter will prove to be important as we determine the quality-adjusted import prices in small markets, in particular, since it potentially takes very efficient firms to overcome the fixed costs of exporting to those markets, and these firms sell high quality. In addition, fixed costs depend on four measures of language similarity between any two countries, F_{ni}^k, $n = 1,..,4$. We consider two random people, one from each country, and construct four probabilities: the probability that they speak a common language; the probability that they speak a language from a common language genus; the probability that they speak a language from a common language family; and the probability that their countries share a common official language. Details are provided in Appendix B.

With these fixed costs, we derive the solution for the gravity equation for X_i^k as:

$$\left(\frac{X_i^k}{M_i(\varphi_i / w_i)^\gamma}\right) = \left(\frac{\overline{p_i^k}}{\left(\kappa_1^k \overline{p_i^{*k}}\right)^{\alpha^k \theta}}\right)^{\frac{-(\sigma-1)(1+\gamma)}{[1+\alpha^k \theta(\sigma-1)]}} \left(tar_i^k \exp\left(\sum_{n=1}^{4} \beta_n F_{ni}^k\right)\right)^{-\left[\frac{\gamma-\alpha^k\theta(\sigma-1)}{1+\alpha^k\theta(\sigma-1)}\right]} \left(\frac{Y^k}{M^k}\right), \qquad (20)$$

where M^k denotes the "market potential" of country k,[10]

[10] This term is the inverse of the "remoteness" variable derived by Chaney (2008). Redding and Venables (2004) refer to (21) as "supplier access" in a monopolistic competition model with homogeneous firms. We are referring to is as "market potential" from the buyer's point of view.

$$M^k \equiv \sum_i M_i \left(\frac{\varphi_i}{w_i}\right)^{\gamma} \left(\overline{p_i^k} \Big/ \left(\kappa_1^k \overline{p_i^{*k}}\right)^{\alpha^k\theta}\right)^{\frac{-(\sigma-1)(1+\gamma)}{[1+\alpha^k\theta(\sigma-1)]}} \left(tar_i^k \exp\left(\sum_{n=1}^{4}\beta_n F_{ni}^k\right)\right)^{-\left[\frac{\gamma-\alpha^k\theta(\sigma-1)}{1+\alpha^k\theta(\sigma-1)}\right]}. \quad (21)$$

The term M^k is higher when there are more firms M_i potentially selling to country *k,* and when transport costs (which affect the c.i.f. and f.o.b. prices (21)) and fixed costs are lower. In practice we obtain M^k as a destination country fixed-effect when estimating the gravity equation. Notice that the real expenditure (Y^k / uv^k) which influences fixed costs does not appear in (20) or (21), because a destination-specific change in fixed costs cancels out: given total expenditure on imports $Y^k = \sum_i X_i^k$, a uniform increase in fixed costs does not affect the application of expenditure across source countries X_i^k. But we will find such a uniform change in fixed costs has an important impact on the quality-adjusted prices.

The corresponding solution for the quality-adjusted prices is:

$$\overline{P_i^k} = \left(\overline{p_i^k} \Big/ \left(\overline{p_i^{*k}} \kappa_1^k \kappa_2^k tar_i^k \exp\left(\sum_{n=1}^{4}\beta_n F_{ni}^k\right)\right)^{\alpha^k\theta}\right)^{\frac{1}{1+\alpha^k\theta(\sigma-1)}} \left(\frac{Y^k}{M^k}\right)^{\frac{\alpha^k\theta}{(1+\gamma)}} \left(\frac{Y^k}{uv^k}\right)^{-\frac{\beta_0\alpha^k\theta}{(1+\gamma)}}. \quad (22)$$

We see that the quality-adjusted price depends on both the c.i.f. and f.o.b. prices as well as the fixed-cost variables (including the *ad valorem* tariffs), and then depends on total import expenditure Y^k in two ways. On the one hand, higher expenditure leads to the entry of less-efficient exporters, who produce lower quality leading to a higher quality-adjusted price. On the other hand, higher real expenditure (Y^k / uv^k) leads to higher fixed costs from (19) so that the marginal exporter must be more efficient, leading to lower quality-adjusted prices. The strength of these two opposing forces depends on the parameter β_0. This parameter is not estimated from the gravity equation so we must look to other data to identify it.

Eaton, Kortum and Kramarz (2011) estimated a regression of the number of firms exporting from France, $\ln N_i^k$, on the log of real manufacturing imports from France across various destination countries, obtaining an elasticity of 0.65.[11] A similar regression on French data is reported by Arkolakis (2010). This regression was repeated in Eaton, Kortum and Sotelo (2012) for Brazil, France, Denmark and Uruguay, yielding an elasticity of 0.71 (or 0.62 with country fixed effects). In a Melitz model with fixed costs specified as in (19), those elasticities measure $(1 - \beta_0)$ so we obtain an estimate for β_0 of about 0.35 for manufacturing as a whole.

In our model, the coefficient of 0.65 linking the number of firms to market size implies an estimate for β_0 of less than 0.35, due to our modeling of fixed costs in (12) as depending on the productivity of the cutoff exporter. To see this, start with (13) where $N_i^k \equiv M_i(\hat{\varphi}_i^k / \varphi_i)^{-\gamma}$ appearing on the right denotes the number (or mass) of exporters. This number is proportional to total exports divided by those of the cutoff exporter, $N_i^k \propto (X_i^k / \hat{X}_i^k)$. Then substitute cutoff exports $\hat{X}_i^k$ from (12) into (13), and simplify to obtain:

$$\left(N_i^k\right)^{(1+\gamma)/\gamma} \propto \left(\frac{X_i^k}{f_i^k}\right) = X_i^k \left(\frac{Y^k}{\overline{uv}^k}\right)^{-\beta_0} \exp\left(\sum_{n=1}^{4} \beta_n F_{ni}^k\right)^{-1},$$

using (19). If the number of exporters has an elasticity of 0.65 with respect to in Y^k, then since exports have elasticity of unity it follows that $0.65\left(\frac{1+\gamma}{\gamma}\right) = 1 - \beta_0$, so that $\beta_0 = 1 - 0.65\left(\frac{1+\gamma}{\gamma}\right)$. Across our industry estimates, γ ranges from about 2 to a very large number, so we see that β_0 will range from about zero up to 0.35. We proceed now with describing the estimation of all other model parameters.

[11] This result is not reported in the published paper, and we thank Jonathan Eaton for informing us of it.

4. Data and Estimation

Data

The primary dataset used is the United Nations' Comtrade database. We compute the bilateral f.o.b. unit values of traded goods using reports from the exporting country. By focusing on the exporters' reports we ensure that these unit values are calculated prior to the inclusion of any costs of shipping the product. The bilateral c.i.f. unit values are calculated similarly using importers' trade reports. Since these unit values include the costs of shipping, we need only add the tariff on the good to produce a tariff-inclusive c.i.f. unit value. To do this we obtain the *ad valorem* tariffs associated with Most Favored Nation status or any preferential status from TRAINS, which we have expanded upon using tariff schedules from the *International Customs Journal* and the texts of preferential trade agreements obtained from the World Trade Organization's website and other online sources. We provide further details in Appendix B.

The independent variation in the importing country's c.i.f. unit value and the exporting country's f.o.b. unit value is essential to identifying their distinct effects in the gravity equation. But it must be admitted that there is a large amount of measurement error in these unit values from the Comtrade database. In fact, it is not unusual for the c.i.f. unit value to be less than the f.o.b. unit value (as can never occur in theory because the former exceeds the latter by transport costs). As an initial step towards correcting for such measurement error, we omitted observations where the ratio of the c.i.f. unit value reported by the importer and the f.o.b. unit value reported by the exporter, for a given 4-digit SITC product and year, was less than 0.1 or exceeded 10. In addition, we omitted such bilateral observations where the c.i.f. value of trade was less than $50,000 in 2005, while adjusting for U.S. inflation so the cutoff was about $25,000 in 1984.

More generally, to reconcile the wide variation in the observed unit values with our model, we assume that the c.i.f. and f.o.b. unit values, denoted by uv_{igt}^{k} and uv_{igt}^{*k}, are related to the true c.i.f. and f.o.b. prices by:

$$\ln uv_{igt}^{k} = \ln(\overline{p_{igt}^{k} / tar_{igt}^{k}}) + u_{igt}^{k}, \text{ and } \ln uv_{igt}^{*k} = \ln \overline{p_{igt}^{*k}} + u_{igt}^{*k}, \tag{23}$$

where u_{igt}^{k} and u_{igt}^{*k} are the measurement errors that are independent of each other and have variances σ_{g}^{k} and σ_{ig}^{*}, respectively. In other words, we are assuming that the measurement error in the c.i.f. unit value for importer *k* does not depend on the source country *i,* while the measurement error in the f.o.b. unit value for exporter *i* does not depend on the importer *k,* and that these errors are independent of each other. We argue in Appendix C that our estimation method is robust to this measurement error in the unit values, which ends up being absorbed by importer and exporter fixed-effects in the estimation. But the errors must be independent for this claim to hold, which is therefore an identifying assumption.

Estimation

To estimate the gravity equation we model the mass of potential exporters in (16) as depending on the estimated labor force L_{igt} involved in the production of exports of good *g* in country *i*, together with country fixed effects:[12]

$$\ln[M_{igt}(\varphi_{igt} / w_{igt})^{\gamma}] = \delta_{g0} \ln L_{igt} + \delta_{ig} + \varepsilon_{igt}^{k}, \tag{24}$$

where ε_{igt}^{k} is a random error. In addition, we make explicit the *ad valorem* tariffs and unit values using in our estimation by re-writing the net-of-tariff c.i.f. price as the import unit value $uv_{igt}^{k} = (\overline{p_{igt}^{k} / tar_{igt}^{k}})$, and the f.o.b. price as the export unit value $uv_{igt}^{*k} = \overline{p_{igt}^{*k}}$. Using these in (20),

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[12] Denoting sectors by *g*, we estimate the labor force in each sector by $L_{it}^{g} \equiv POP_{it}(X_{it}^{g} / GDP_{it})$, or country population times exports in sector *g* divided by country GDP.

we readily obtain exports expressed as the difference between countries *i* and *j* selling to destination *k*:

$$\ln X_{igt}^{k} - \ln X_{jgt}^{k} = -A_g^k \left[\left(\ln \overline{uv_{igt}^{k}} - \ln \overline{uv_{jgt}^{k}} \right) - \alpha_g^k \theta_g \left(\ln \overline{uv_{igt}^{*k}} - \ln \overline{uv_{jgt}^{*k}} \right) \right] + \delta_{0g} (\ln L_{git} - \ln L_{gjt})$$
$$+ \delta_{ig} - \delta_{jg} - \sum_{n=1}^{4} B_g^k \beta_{ng} (F_{ni}^k - F_{nj}^k) - C_g^k \left(\ln tar_{igt}^k - \ln tar_{igt}^k \right) + \varepsilon_{igt}^k - \varepsilon_{jgt}^k, \quad (25)$$

where:

$$A_g^k \equiv \frac{(\sigma_g - 1)(1+\gamma_g)}{1+\alpha_g^k \theta_g (\sigma_g - 1)},\ B_g^k \equiv \frac{\gamma_g - \alpha_g^k \theta_g (\sigma_g - 1)}{1+\alpha_g^k \theta_g (\sigma_g - 1)},\ \mathrm{C}_g^k \equiv \beta_{5g} \left[\frac{\sigma_g (1+\gamma_g)}{1+\alpha_g^k \theta_g (\sigma_g - 1)} \right] - 1. \quad (26)$$

Three features of this estimating equation deserve attention. First, notice that the c.i.f. unit values appear with the negative coefficient $-A_g^k$ in this gravity equation, whereas the f.o.b. unit values appear with a *positive* coefficient of $A_g^k \alpha_g^k \theta_g$, because they represent product quality and higher quality leads to great demand.[13] The key to successful estimation will be to obtain this sign pattern on the c.i.f. and f.o.b. unit values, which we achieve by adapting the GMM estimator from Feenstra (1994). As described in Appendix C, we add a simple supply specification whereby the specific and iceberg trade costs depend on distance, *ad valorem* tariffs and the quantity traded. Feenstra (1994) assumed that the supply shocks and demand shocks are uncorrelated. That assumption seems unlikely to hold with unobserved quality, since a change in quality could shift both supply and demand. But here, the demand errors and the supply errors are the residuals after *taking into account* quality. So the assumption that they are uncorrelated seems much more acceptable, and is the basis for the GMM estimation.

Second, we have isolated one plus the *ad valorem* tariffs tar_{ig}^k in (25)-(26) because this

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[13] For the purpose of estimating (19), we simplify the analysis slightly by taking the time-average of the ratios of real GDP in (9b), so that α_g^k depends on the SITC good *g* but not on time.

variable plays a special role in the estimation. In theory, the term β_{5g} appearing within C_g^k is unity, so that $C_g^k = A_g^k + B_g^k$, meaning that *ad valorem* tariffs have an impact through their effect on the consumer price and on fixed costs. But in the estimation we allow for $\beta_{5g} \neq 1$ so that tariffs can have a distinct impact on export flows. We hypothesize that tariff evasion can lead to $\beta_{5g} < 1$, so that tariffs have a smaller impact in the data than indicated by the theory.

The final challenge is that not all the parameter estimates are identified without additional information. In particular, we estimate $B_g^k \beta_{ng}$ in (25) but not these coefficients alone. If we do not identify B_g^k, then we cannot solve for σ_g and γ_g. We resolve this issue as in Chaney (2008), by using estimates of $\zeta_g^{US} = \gamma_g / [\alpha_g^{US} \theta_g (\sigma_g - 1)]$ from regressions of firm rank on size for each SITC sector in the U.S., where we further normalize $\alpha_g^{US} \equiv 1$.[14] Then for other countries, $\zeta_g^k = \gamma_g / [\alpha_g^k \theta_g (\sigma_g - 1)] \Leftrightarrow \zeta_g^k \alpha_g^k = \gamma_g / [\theta_g (\sigma_g - 1)] = \zeta_g^{US} \alpha_g^{US} = \zeta_g^{US}$. It follows that γ_g is obtained as $\zeta_g^{US} \theta_g (\sigma_g - 1)$. Substituting this into (25)-(26), we obtain an estimating equation that is nonlinear in the parameters σ_g, θ_g and λ_g.

In addition, it is difficult to identify λ_g from the gravity equation (25) alone. For this reason, we rely instead on the estimates of λ_g that come from the f.o.b. price regressions reported in section 2. In those regressions, we use preliminary estimates of σ_g and θ_g that come from estimating (25) and (9b).[15] Using these preliminary estimates, we estimate (9a) and (9b) to obtain improved values for λ_g, as reported in section 2. These improved values are substituted into (25) to obtain new estimates for σ_g and θ_g.

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[14] We thank Thomas Chaney for providing these estimates for 3-digit SITC Rev. 3 sectors for the United States, which we concorded to 3-digit SITC Rev. 2 sectors. The normalization $\alpha_g^{US} \equiv 1$ is harmless because α_g^k always appears multiplied by θ, so $\alpha_g^{US} \equiv 1$ fixes the value for θ in our estimates.

[15] In those preliminary estimates we constrain λ_g to be non-negative by estimating it as $\lambda_g = \mu_g^2$, and find that the median estimate is 0.005 with about one-third of the estimates across the SITC industries estimated at zero.

5. Parameter Estimates

Estimation is performed for each 4-digit SITC Revision 2 good (which we also refer to as an industry) using bilateral trade between all available country pairs during 1984-2008. There are 2.4 million observations with data on both the c.i.f. and f.o.b. unit values, and excluding those goods with fewer than 50 observations, we perform the GMM estimation on 783 industries as shown in the first row of Table 1.[16] The median estimate of σ_g is 6.39, not counting seven industries with an inadmissible value less than unity; the median estimate of γ_g is 8.85, not counting the same seven industries with an inadmissible negative value; and the median estimate of θ_g is 0.64, not counting one case with an inadmissible value greater than unity. For inadmissible values or for SITC industries with fewer than 50 observations, we replace the parameter estimates with the median estimate from the same 3-digit or 2-digit SITC industry, after which we find the median estimates shown in the last row of Table 1 for 925 industries.

The frequency distribution of parameter estimates are illustrated in Figures 3-5. Our median estimate for the elasticity of substitution σ_g is higher than found by Broda and Weinstein (2006) for the United States. We have found that our higher value comes from using worldwide trade data and correcting for quality, and from using an empirical specification that is more robust to measurement error since we do not take differences over time and instead include source-country fixed effects in our estimation of (25).[17] Our median estimate for the Pareto parameter γ is quite close to that reported by Eaton and Kortum (2002), who also considered bilateral trade between many countries.[18]

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[16] In each industry we use only the most common unit of measurement, which is nearly always kilograms.

[17] Destination country fixed effects are implicitly included, too, because (19) is specified as the difference between countries *i* and *j* exporting to country *k*.

[18] This median estimate is higher, however, than the recent results of Simonovska and Waugh (2011, 2012).

Table 1: Median Parameter Estimates

GMM Estimation Method with:	Number of SITC industries	σ	γ	θ
Dropping SITC4 with < 50 observations	783	6.39	7.95	0.63
No. of inadmissible parameters	8	7	7	1
Filling in SITC4 with < 50 observations or inadmissible parameters	925	6.43	8.85	0.64

Figure 3: Frequency Distribution for Estimates of σ_g
(Note: Estimates are right-censored for presentation purposes only)

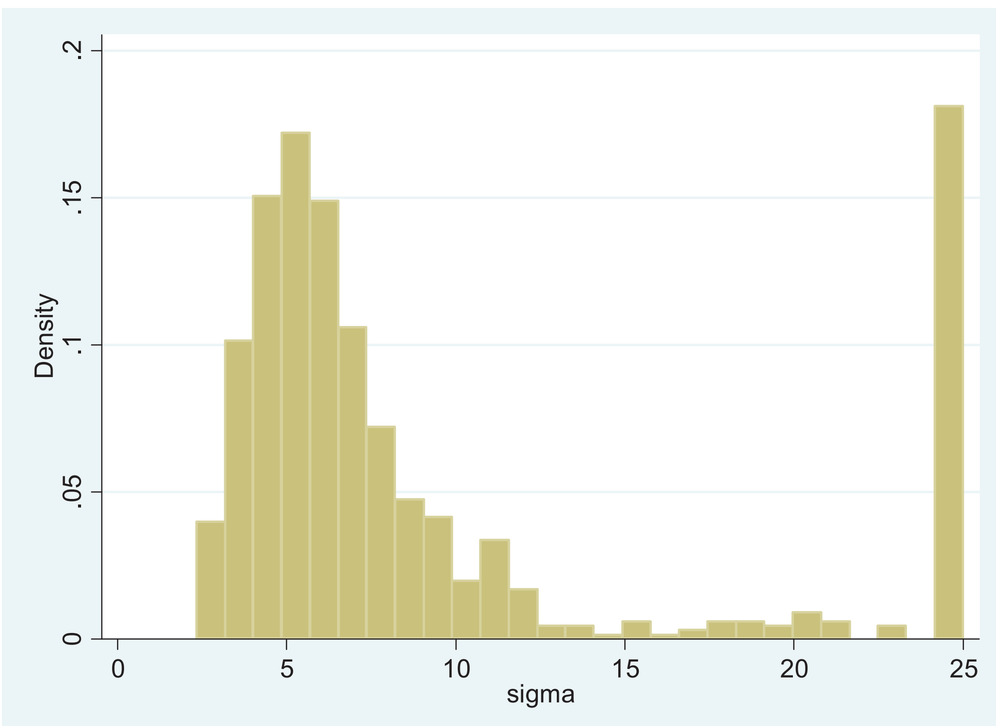


Figure 4: Frequency Distribution for Estimates of γ_g
(Note: Estimates are right-censored for presentation purposes only)

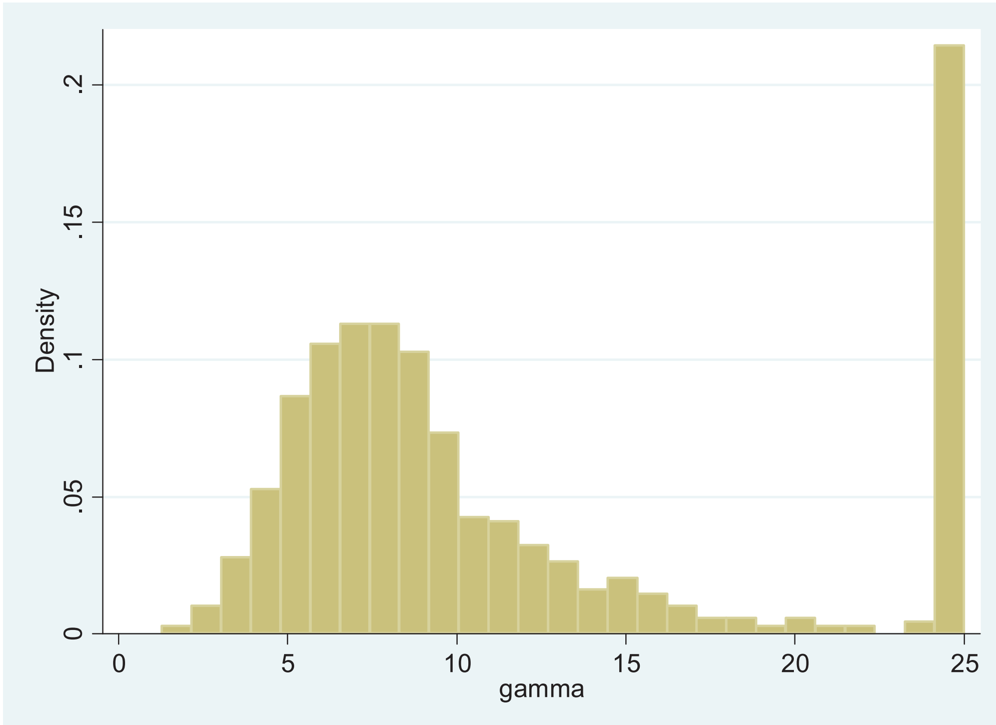
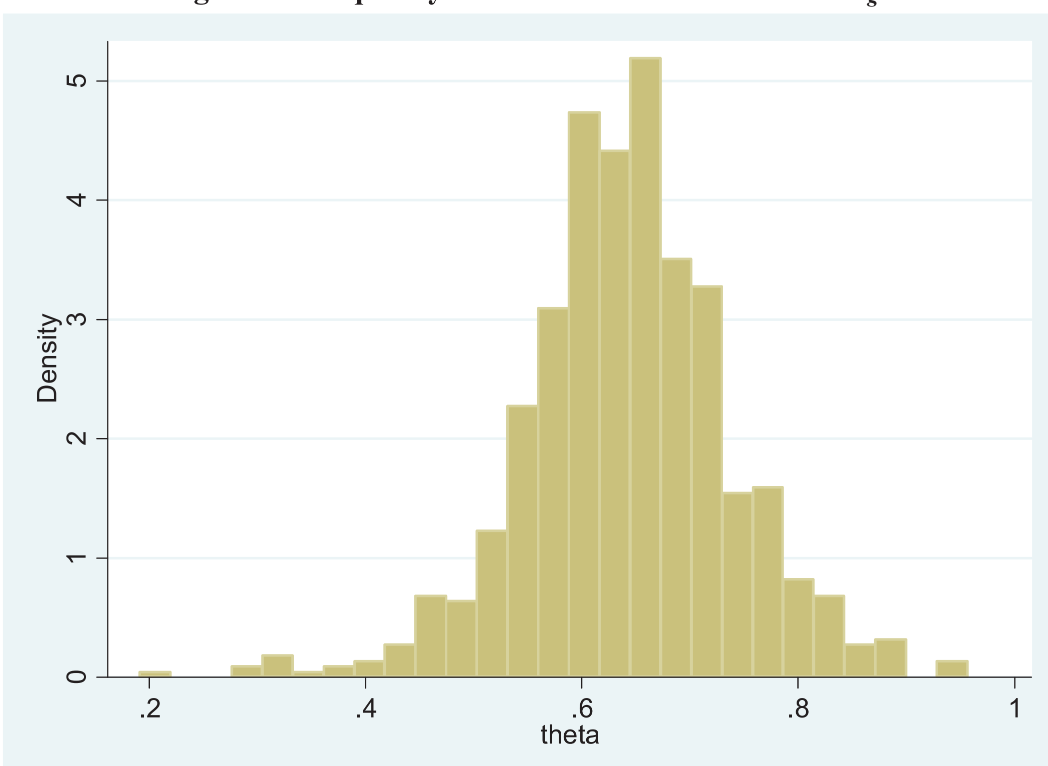


Figure 5: Frequency Distribution for Estimates of θ_g



Turning to θ_g, Crozet, Head and Meyer (2012) are the only other authors to estimate the elasticity of quality with respect to inputs. They use ratings of wines to obtain an estimate of 0.29 for Champagne, which is at the lower end of our estimates shown in Figure 3. However, our estimates for θ_g depend on the normalization we have adopted for α_g^k, namely, that $\alpha_g^{US} \equiv 1$. If we had instead normalized $\alpha_g^k \equiv 1$ for a lower-income country so that $\alpha_g^{US} > 1$, then all of our estimates of θ_g would be correspondingly lower.

6. Indexes of Quality-Adjusted Prices and Quality

Quality-adjusted prices are shown by (22). When comparing two exporters *i* and *j* selling to the same destination *k*, we write that expression while adding the good and time subscripts as,

$$\frac{\overline{\overline{P_{igt}^k}}}{\overline{\overline{P_{jgt}^k}}} = \left(\frac{\overline{p_{igt}^k} / \overline{p_{igt}^{*k}}^{\alpha_g^k \theta_g}}{\overline{p_{jgt}^k} / \overline{p_{jgt}^{*k}}^{\alpha_g^k \theta_g}} \right)^{\frac{1}{[1+\alpha_g^k \theta_g (\sigma_g - 1)]}} \left(\frac{tar_{igt}^k \exp\left(\sum_{n=1}^4 \beta_{ngt} F_{ni}^k\right)}{tar_{jgt}^k \exp\left(\sum_{n=1}^4 \beta_{ngt} F_{nj}^k\right)} \right)^{\frac{-\alpha_g^k \theta_g}{[1+\alpha_g^k \theta_g (\sigma_g - 1)]}} . \quad (27)$$

Notice that the market potential term M_{gt}^k appearing in (22) cancels out when computing the ratio, which means that this term will not enter our calculation of relative quality-adjusted *export* prices. Likewise, the market size term (Y_{gt}^k / uv_{gt}^k) which influences fixed costs from (19) does not appear and the terms κ_{1g}^k and κ_{2g}^k do not appear, since both exporters are selling to the same destination market.

The calculation of quality-adjusted prices on the import side is slightly more complicated. If we compare a given exporter *i* selling to two destinations *k* and *l*, then the ratio of (22) would involve two different taste parameters α_g^k and α_g^l, and two market potential terms M_{gt}^k and

M_{gt}^{l}. The latter are estimated as destination-country fixed effects in the gravity equation (see Appendix C), and are particularly prone to absorbing measurement error in any of the unit-values included there. We can avoid this problem by instead using (16) to evaluate the quality-adjusted prices, where the term $M_{igt}(\varphi_{igt} / w_{igt})^{\gamma}$ appearing there cancels out in the ratio. The problem of having two different taste parameters α_g^k and α_g^l is resolved by using an average value $\bar{\alpha}_g$ for all countries importing the good g.[19] Then we measure the ratio of (16) for a given exporter i selling to two destinations k and l as:

$$\frac{\overline{P_{igt}^{k}}}{\overline{P_{igt}^{l}}} = \left(\frac{\overline{p_{igt}^{k}} / \left(\kappa_{1g}^{k} \overline{p_{igt}^{*k}} \right)^{\bar{\alpha}_g \theta_g}}{\overline{p_{igt}^{l}} / \left(\kappa_{1g}^{l} \overline{p_{igt}^{*l}} \right)^{\bar{\alpha}_g \theta_g}} \right) \left(\frac{X_{igt}^{k} / \kappa_{2g}^{k} tar_{igt}^{k} \left(\dfrac{Y_{gt}^{k}}{uv_{gt}^{k}} \right)^{\beta_{0g}} \exp\left(\sum_{n=1}^{4} \beta_{ngt} F_{ni}^{k} \right)}{X_{igt}^{l} / \kappa_{2g}^{l} tar_{igt}^{l} \left(\dfrac{Y_{gt}^{l}}{uv_{gt}^{l}} \right)^{\beta_{0g}} \exp\left(\sum_{n=1}^{4} \beta_{ngt} F_{ni}^{l} \right)} \right)^{\frac{\bar{\alpha}_g \theta_g}{(1+\gamma_g)}} \tag{28}$$

To implement (27) and (28), we use the import c.i.f. unit value inclusive of one plus the *ad valorem* tariff, $uv_{igt}^{k} tar_{igt}^{k}$, to replace $\overline{p_{igt}^{k}}$, and the f.o.b. unit value uv_{igt}^{*k} to replace $\overline{p_{igt}^{*k}}$. Also, we collect terms involving the tariff variable tar_{igt}^{k}, which appears within the tariff-inclusive prices $\overline{p_{igt}^{k}}$ and multiplying the fixed costs. Just as in our estimation of the gravity equation (25)-(26), we allow the coefficient on the tariff to differ from its theoretical value using the estimated coefficient β_{5g}. For consistency with the import side we also used the average $\bar{\alpha}_g$ for the export side in (27), though this has a minimal impact on the export results.

[19] According to Fisher and Shell (1972), with changing preferences (in this case changing between countries), a suitable approach is to compute a geometric mean of price indexes that first uses one country's preferences and then uses the other's. We have not (yet) implemented the Fisher-Shell approach for our indexes, but instead we evaluate quality-adjusted prices in (27)-(28) using an average preference for quality $\bar{\alpha}_g$.

The c.i.f. unit values uv_{igt}^{k} and the quality-adjusted unit-values denoted by UV_{igt}^{k}, obtained from (27)-(28), are then aggregated from the 4-digit SITC to the Broad Economic Categories (BEC) to obtain overall indexes of quality and quality-adjusted prices of exports and imports for each country and year in our dataset. The formula we shall use for aggregation is the so-called GEKS method,[20] which is a many-country generalization of Fisher Ideal indexes, as we shall describe. What we add to this method is a two-stage aggregation procedure that arises naturally from our trade data.

Indexes for Export Prices and Quality

In the first stage, for each 4-digit SITC product *g* we aggregate over all partner countries in trade, i.e. over all destination countries for an exporter and all source countries for an importer. Consider first the problem from the exporters' point of view. The unit-value ratio $(uv_{git}^{k} / uv_{gjt}^{k})$ compares countries *i* and *j* selling to *k,* from we shall construct an index of *relative export prices.* That is, we compare the unit values of countries *i* and *j* only when they are selling to the same country *k:* essentially, we are treating products sold to different countries as entirely different goods and avoid comparing their prices in that case.

Suppose that exporting countries *i* and *j* both sell the 4-digit SITC product *g* to k=1,…,C_{ij} destination markets. The Laspeyres and Paasche price indexes of these export unit values are:

$$P_{gijt}^{L} \equiv \frac{\sum_{k=1}^{C_{ij}} uv_{git}^{k} q_{gjt}^{k}}{\sum_{k=1}^{C_{ij}} uv_{gjt}^{k} q_{gjt}^{k}}, \quad \text{and,} \quad P_{gijt}^{A} \equiv \frac{\sum_{k=1}^{C_{ij}} uv_{git}^{k} q_{git}^{k}}{\sum_{k=1}^{C_{ij}} uv_{gjt}^{k} q_{git}^{k}}. \tag{29}$$

[20] Named after Gini, Eltetö and Köves, and Szulc. We refer the reader to Balk (2008) and Deaton and Heston (2010) for a modern treatment and details of these historical references. While the Fisher Ideal index is not "exact" for a CES utility function, it belongs to the class of superlative indexes, and Diewert (1978) argues that these indexes approximate each other quite closely. We employ it here because it is commonly used by statistical agencies, including the ICP and PWT, which also use the GEKS generalization.

In these expressions, q^k_{git} and q^k_{gjt} are the quantity exported by countries *i* and *j* to country *k.* Alternatively, we could instead use the *quality-adjusted* unit values UV^k_{git} in these formulas, in which case the quantities are instead Q^k_{git} with $uv^k_{git}q^k_{git} = UV^k_{git}Q^k_{git}$ and likewise for country *j,* so the export *values* are not affected by the quality adjustment. Regardless of whether the unit values or quality-adjusted unit values are used, the Laspeyres and Paasche index can always be re-written as a weighted average of their ratios. Letting $s^k_{gjt} = uv^k_{gjt}q^k_{gjt} / \sum_k uv^k_{gjt}q^k_{gjt}$ denote the export shares for country *j,* the Laspeyres index in (29) equals $P^L_{gijt} = \sum_k s^k_{gjt}(uv^k_{git} / uv^k_{gjt})$. Likewise, the Paasche index is a weighted average of the unit-value ratios using the export shares s^k_{git} of country *i.* In either case, we can alternatively use the ratio of quality-adjusted unit values, $(UV^k_{git} / UV^k_{gjt})$. In this way, we obtain the Laspeyres and Paasche indexes for both unit values and quality-adjusted unit values.

The Fisher Ideal price index is the geometric mean of the Laspeyres and Paasche indexes, $P^F_{gijt} = (P^L_{gijt}P^A_{gijt})^{0.5}$. Then the GEKS price index of country *i* relative to *k* is computed by taking the mean over all Fisher indexes for exports of country *i* relative to exports of *j* times the Fisher index for exports of *j* relative to exports of *k*:

$$P^{GEKS}_{gikt} \equiv \prod_{j=1}^{C} \left(P^F_{gijt}P^F_{gjkt}\right)^{1/C}, \tag{30}$$

with $P^F_{giit} \equiv 1$ for $i = 1,\ldots,C$. In most applications, the resulting GEKS indexes are transitive.[21] That property does not necessarily hold in our case, however, because two countries may not export the 4-digit SITC product to the same set of partners, so that the mean in (30) is actually taken over only

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[21] This is shown from (29) by noting that $P^{GEKS}_{jkt} = 1 / P^{GEKS}_{kjt}$, so that we readily compute $P^{GEKS}_{ikt}P^{GEKS}_{kmt} = P^{GEKS}_{imt}$.

the set of exporters *j* that share some common destination markets with both countries *i* and *k*. Despite the fact that transitivity may not hold, the GEKS transformation of the Fisher Ideal indexes in (30) is useful because it compares the export prices of countries *i* and *k* (selling to the same destination markets) via all possible indirect comparisons with other exporters.[22]

This GEKS aggregation is done for each 4-digit SITC product. We trim one percent of the estimated quality-adjusted price indexes (i.e. the upper and lower 0.5 percent) and then proceed with the second stage aggregation over the SITC products *g*. We again use Fisher Ideal indexes – now computed by summing over products rather than over partner countries as in (29) – together with the GEKS transformation. In this second step we choose the United States as the comparison country *k*, so we end up with indexes of unit values, or quality-adjusted unit-values, for each exporting country and year relative to the United States. These indexes are computed for all exports and for the one-digit Broad Economic Categories (BEC). The BEC distinguishes food and beverages, other consumer goods, capital, fuels, intermediate inputs, and transport equipment, so this breakdown of sectors should be useful for many researchers interested in international prices.

We refer to the GEKS index of unit values as the "price index" and the GEKS index of quality-adjusted unit values as the "quality-adjusted price index". Our final step is to divide the former by the latter – for each country, year and BEC– to obtain the index of export quality.[23]

Indexes for Import Prices and Quality

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[22] To maximize the number of indirect comparisons, for each 4-digit SITC product and year we chose the base country *k* as the exporter having the largest number of destination markets times its total exports to all of them.
[23] Since our indexes of prices and quality-adjusted prices are both measured as ratios relative to the United States, the same is true of quality. For expositional convenience in our table and graphs, we re-normalize these series so they are measured as ratios relative to their (geometric) means, so the world average price, quality-adjusted price, and quality are all unity. When the quality-adjusted prices are incorporated into the next generation of PWT, then time-series variation is introduced by re-normalizing them again so the U.S. series follows the national accounts prices for U.S. import and exports at the BEC 1-digit level (see Feenstra, Inklaar and Timmer, 2012)."

Our treatment of imports is similar to our treatment of exports, so we only highlight the differences. In the first stage, the Laspeyres and Paasche indexes are computed by summing over source countries *i* that importers *k* and *l* both purchase from. So we compare the import prices of countries *k* and *l* only if they come from the same exporter *i.* As we found earlier, the Laspeyres and Paasche indexes can be expressed as share-weighted averages of the unit-value ratio, or quality-adjusted unit-value ratio, for countries *k* relative to *l.* That is, the Laspeyres and Paasche indexes depend on $uv_{igt}^{k} / uv_{igt}^{l}$, or alternatively on the quality-adjusted unit value $UV_{igt}^{k} / UV_{igt}^{l}$. We then compute the Fisher Ideal indexes and perform the GEKS transformation, resulting in an index of the import prices for country *k* relative to a base country *m* for each SITC product.[24] In the second stage, we aggregate over products *g* to obtain indexes of import prices, and quality-adjusted prices, relative to the United States for each BEC category. Dividing the former by the latter, we obtain the import index of quality.

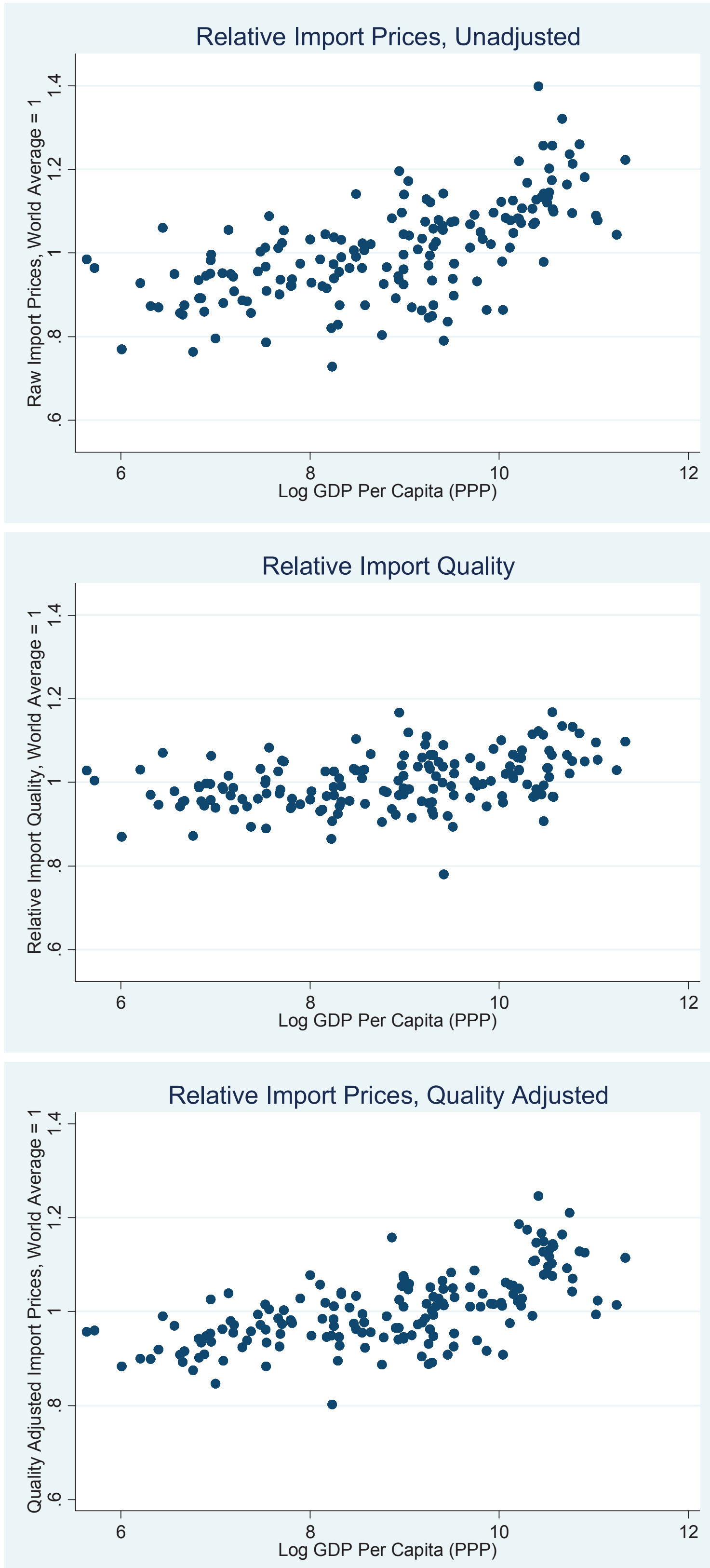
Empirical Results

Figures 6 to 8 summarize our results for 2007. We aggregate raw export prices for each 4-digit SITC product into an aggregate price index for exports, with the average export price index across countries normalized to 1. We then similarly aggregate our quality estimates, and plot these indexes for 2007 for about 200 countries in Figure 6. The results broadly conform with our priors. Developed countries tend to export more expensive goods (top panel), and we estimate these goods to be of higher than average quality (second panel). The quality adjusted-price (price divided by quality), about which we have less strong priors, tends to be only slightly higher for developed countries (bottom panel).

[24] Analogous to the export side, for each 4-digit SITC product and year we chose the base country *l* as the importer having the largest number of source countries times its total imports from all of them.

Figure 6: Exports - Raw Prices, Quality, and Quality Adjusted Prices in 2007

Figure 7: Relative Import Unit Values, Quality, and Quality Adjusted Prices in 2007



We illustrate a similar exercise for import prices in Figure 7. Developed countries import more expensive items (top panel) that are of higher quality (second panel). Quality-adjusted import prices (third panel) increase noticeably with the importing country's GDP per capita. This pattern is due to an interaction of preferences for quality and the rising marginal cost of producing quality. Rich countries tend to prefer higher quality goods – this is reflected in our estimates of λ_g in equation (9a) – which enter the import quality-adjusted price in (28) via κ_{1g}^k and κ_{2g}^k . But our estimates of θ_g between zero and unity means, from (4), that there is an amplified effect of quality on increasing the marginal cost, so that higher quality induced by a preference for quality leads to a higher quality-adjusted price.

It is evident the variation in quality-adjusted import prices in Figure 7 is much greater than for export prices in Figure 6. Numerically, this occurs because the preference for quality affects import prices in (28), along with bilateral imports X_{igt}^k and total import expenditure Y_{gt}^k, none of which enter the export-side formula in (27). The economic intuition for this result comes because relative import prices are obtained by comparing a given exporter *i* selling to two destinations *k* and *l*. In our model, any difference in the f.o.b. price from a given exporting firm must be due to quality. As we noted earlier in (10), log quality is only a fraction of the log f.o.b. price, with the remaining difference in f.o.b. prices in (11) attributed to the quality-adjusted price. This pattern is illustrated on the import side in Figure 7.

For exports, we instead compare two countries *i* and *j* selling to a given destination market *k*, and the intuition is different. Now the differences in the exporter firms f.o.b. prices can be due to costs or quality. But because they are selling to the same market, and our estimated elasticities of substitution σ are quite high, there is little scope for average quality-adjusted prices to differ by exporter, at least when there are meaningful amounts of exports. Hence, most of the difference in price is attributed to quality, as shown in Figure 6.

Figure 8: Terms of Trade 2007 - Unadjusted and Quality Adjusted

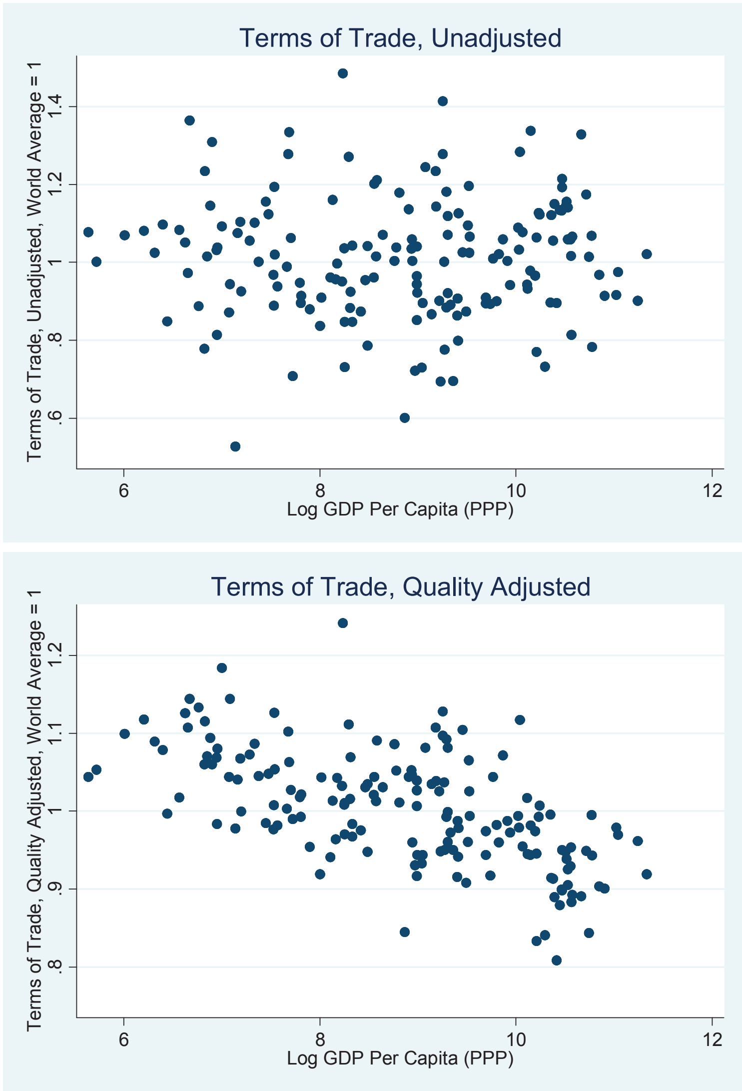


Figure 8 shows terms of trade estimates for 2007. Terms of trade estimates constructed using raw export and import prices fluctuate substantially across countries, and lie between 0.52 and 1.49. Terms of trade estimates constructed from quality-adjusted prices move in a much narrower band, between 0.80 and 1.24. Despite the narrowness of this band, these quality-adjusted terms of trade measures are sufficiently different from unity to produce meaningful differences between output-based and expenditure-based real GDP estimates for many countries in the PWT (Feenstra, Inklaar and Timmer, 2012). Notably, the terms of trade decline in real

GDP per capita, as wealthier countries are trading higher-quality goods at higher quality-adjusted prices, but this effect is much stronger for imports than for exports.

We report estimates for aggregate export quality for 1987, 1997 and 2007 in Table 2 for the 52 largest traders measured by their average value of exports from 1984 to 2008. Swiss exports have the highest quality, on average 68% higher than the world average in 2007, followed by Finland and Israel with quality 39 and 37 percent higher, respectively, than the average. Japan, the U.S. and other wealthy European countries usually have 15 to 30 percent higher export quality than the average. Of note are the recent quality increases for several Eastern European countries that have joined the EU, especially those proximate to Germany: Czech Republic, Hungary, Poland and Slovakia. Asian countries that rapidly industrialized in the 1970's – Hong Kong, Singapore, South Korea and Taiwan – exhibit improving export quality but with the exception of Singapore still lag average export quality. Poorer large Asian countries have notably lower quality, with Indian and Chinese export quality respectively 13 percent and 34 percent lower than average levels. Vietnam and Indonesia do little better, with quality lagging average levels in 2007 by 13 and 21 percent respectively.

It is interesting that China's relative export quality appears to have declined despite substantial economic progress. This does not imply that its *absolute* export quality has declined, since other countries may have raised quality. China’s substantial exports of relatively low-quality products may have in fact caused most other countries to focus on higher quality goods; see Amiti and Khandelwal (2009) for a discussion. We can find plenty of examples in the detailed data of rising relative quality for China, such as “Computers” rising from 0.27 in 1987 to 0.35 in 1997 and 0.67 in 2007, or “Coarse Ceramic Housewares” (dinnerware), rising from 0.36 in 1987 and 1997 to 0.51 in 2007, or “Footwear”, rising from 0.22 in 1987 to 0.49 in 1997 and

Table 2: Export Quality in 1987, 1997 and 2007.
Quality Rankings, All Goods

	Rank				Normalized Quality, World Average = 1			
country	1987	1997	2007	Change	1987	1997	2007	Change
Switzerland	1	1	1	0	1.53	1.58	1.68	0.15
Finland	3	5	2	1	1.24	1.32	1.39	0.15
Israel	16	4	3	13	1.15	1.37	1.37	0.21
Austria	2	9	4	-2	1.28	1.24	1.32	0.04
United Kingdom	15	7	5	10	1.16	1.31	1.31	0.15
Ireland	9	6	6	3	1.19	1.32	1.31	0.12
Sweden	7	2	7	0	1.21	1.38	1.27	0.06
France	5	13	8	-3	1.23	1.20	1.26	0.04
Germany*	4	10	9	-5	1.23	1.22	1.25	0.01
Denmark	8	8	10	-2	1.20	1.27	1.24	0.04
Japan	12	3	11	1	1.16	1.37	1.24	0.07
USA	6	15	12	-6	1.22	1.19	1.22	0.00
Australia	19	12	13	6	1.09	1.20	1.21	0.11
New Zealand	18	14	14	4	1.15	1.20	1.19	0.04
Norway	11	18	15	-4	1.16	1.16	1.19	0.03
Canada	25	16	16	9	1.01	1.16	1.18	0.17
Italy	13	11	17	-4	1.16	1.22	1.18	0.02
Belgium	17	19	18	-1	1.15	1.13	1.15	0.01
Netherlands	21	20	19	2	1.08	1.11	1.14	0.07
Portugal	22	25	20	2	1.07	1.06	1.14	0.07
Nigeria	33	26	21	12	0.94	1.05	1.12	0.18
Chile	29	22	22	7	0.96	1.11	1.11	0.15
Spain	30	17	23	7	0.96	1.16	1.11	0.15
Algeria	23	31	24	-1	1.04	1.00	1.10	0.06
Singapore	27	21	25	2	0.98	1.11	1.06	0.08
Hungary	46	35	26	20	0.81	0.95	1.05	0.25
South Africa	36	30	27	9	0.91	1.01	1.03	0.13
Saudi Arabia	20	32	28	-8	1.09	1.00	1.01	-0.07
Slovakia*	50	48	29	21	0.70	0.82	1.01	0.31
Mexico	38	43	30	8	0.89	0.87	1.01	0.12
Czech Rep.*	51	42	31	20	0.70	0.88	1.00	0.30
Colombia	24	24	32	-8	1.02	1.07	0.97	-0.04
Russia*	43	40	33	10	0.84	0.88	0.97	0.14
UAE	10	23	34	-24	1.17	1.07	0.97	-0.21
Turkey	31	33	35	-4	0.96	0.99	0.95	-0.01
Argentina	40	27	36	4	0.88	1.04	0.95	0.07
Rep. of Korea	37	29	37	0	0.90	1.01	0.94	0.05
Philippines	42	37	38	4	0.86	0.92	0.94	0.08
Romania	48	50	39	9	0.79	0.78	0.94	0.14
Iran	14	34	40	-26	1.16	0.97	0.94	-0.22
Brazil	28	28	41	-13	0.97	1.02	0.93	-0.04
Poland	52	46	42	10	0.68	0.85	0.92	0.24
Thailand	45	41	43	2	0.83	0.88	0.90	0.07
Malaysia	35	45	44	-9	0.91	0.86	0.89	-0.02
Hong Kong	41	49	45	-4	0.87	0.81	0.88	0.01
Viet Nam	39	36	46	-7	0.88	0.94	0.87	-0.01
India	34	39	47	-13	0.92	0.91	0.87	-0.05
Venezuela	32	44	48	-16	0.96	0.86	0.86	-0.09
Taiwan	47	47	49	-2	0.79	0.84	0.86	0.07
Ukraine*	44	51	50	-6	0.84	0.74	0.80	-0.03
Indonesia	26	38	51	-25	0.99	0.92	0.79	-0.20
China	49	52	52	-3	0.76	0.69	0.66	-0.10
Mean:					1.01	1.06	1.07	
Standard Deviation:					0.18	0.19	0.19	

(* 1987 data are from West Germany, Czechoslovakia, Czechoslovakia, USSR and USSR respectively)

0.84 in 2007. But there are an almost equal number of examples of falling relative quality. At the SITC 4-digit level the median quality estimate for China has risen modestly from 0.55 in 1987 to 0.57 in 1997 and 0.61 in 2007. What is working against China in aggregate are the weights applied to items due to compositional shifts in China's exports. In 1987, 62 percent of China's exports were in BEC categories 1 through 3: Food, Industrial Supplies, and Fuels. China's measured quality was much closer to average levels for these products, varying from 0.85 for Food to 0.99 for Fuels. By 1997 these exports had declined to 35 percent of China's exports, and to just 27 percent by 2007. China's exports at first were mostly re-oriented towards consumer goods (BEC 6), with that share rising from 30 percent in 1987 to 44 percent in 1997, but these declined back to 27 percent in 2007. The more prolonged re-orientation was towards capital goods and parts (BEC 4), rising from 3 percent of China's exports in 1987 to 17 percent in 1997 and 39 percent in 2007. It is in capital goods and parts where China's relative export quality has always been lowest, between 36 and 52 percent of average levels. China's re-allocation from sectors of relatively high quality towards sectors with relatively low quality is also helping to mask the quality improvements that we often observe as consumers.

Tables 3 through 8 report export quality results for the top-20 exporters in each 1-digit Broad Economic Category (BEC). With a few notable exceptions, the pattern for aggregate quality holds in each of the BEC categories: rich countries tend to have high quality in all BEC categories, while poor countries tend to have notably lower quality. The main exceptions are in Table 5 for BEC 3: Fuels and Lubricants, where there is a less clear relationship between export quality and the exporter's level of development. The recent improvement in Eastern European quality is very apparent in their transport equipment exports. China's declining aggregate relative quality also appears in BEC 1: Food and Beverages and BEC 2: Industrial Supplies. China's

Table 3: Export Quality in 1987, 1997 and 2007.
Quality Rankings, BEC 1: Food and Beverages

	Rank				Normalized Quality, World Average = 1			
country	1987	1997	2007	Change	1987	1997	2007	Change
France	3	1	1	2	1.20	1.26	1.41	0.21
United Kingdom	4	2	2	2	1.18	1.25	1.36	0.18
Australia	10	4	3	7	1.09	1.20	1.36	0.27
Ireland	2	5	4	-2	1.20	1.20	1.34	0.14
Italy	8	7	5	3	1.14	1.13	1.23	0.09
Denmark	1	3	6	-5	1.24	1.21	1.21	-0.03
Germany*	5	11	7	-2	1.17	1.10	1.19	0.01
Netherlands	11	15	8	3	1.07	1.06	1.18	0.11
Belgium	7	8	9	-2	1.14	1.11	1.17	0.03
USA	9	6	10	-1	1.12	1.14	1.16	0.04
New Zealand	6	10	11	-5	1.16	1.10	1.15	-0.01
Spain	15	12	12	3	0.94	1.08	1.12	0.18
Canada	14	9	13	1	1.01	1.10	1.11	0.10
Malaysia	16	13	14	2	0.91	1.07	1.01	0.10
Mexico	17	19	15	2	0.89	0.88	1.01	0.11
Brazil	13	17	16	-3	1.02	1.03	0.92	-0.10
Thailand	20	16	17	3	0.77	1.04	0.90	0.13
Indonesia	12	14	18	-6	1.02	1.07	0.85	-0.17
Argentina	19	20	19	0	0.80	0.87	0.85	0.06
China	18	18	20	-2	0.85	0.95	0.82	-0.03
Mean:					1.05	1.09	1.12	
Standard Deviation:					0.14	0.11	0.18	

(* 1987 data are from West Germany)

Table 4: Export Quality in 1987, 1997 and 2007.
Quality Rankings, BEC 2: Industrial Supplies

	Rank				Normalized Quality, World Average = 1			
country	1987	1997	2007	Change	1987	1997	2007	Change
Switzerland	1	2	1	0	1.61	1.53	1.52	-0.09
Japan	2	1	2	0	1.61	1.71	1.51	-0.09
United Kingdom	5	3	3	2	1.28	1.37	1.37	0.09
France	7	8	4	3	1.25	1.21	1.28	0.02
Sweden	3	4	5	-2	1.32	1.30	1.27	-0.05
USA	9	7	6	3	1.21	1.22	1.27	0.06
Austria	4	10	7	-3	1.31	1.20	1.27	-0.04
Germany*	8	11	8	0	1.25	1.18	1.26	0.02
Italy	6	5	9	-3	1.26	1.25	1.22	-0.03
Hong Kong	13	6	10	3	1.11	1.25	1.22	0.11
Netherlands	10	15	11	-1	1.14	1.08	1.18	0.04
Spain	16	16	12	4	1.02	1.07	1.16	0.14
Canada	17	12	13	4	1.00	1.14	1.14	0.14
Australia	11	14	14	-3	1.13	1.09	1.14	0.01
Rep. of Korea	12	9	15	-3	1.12	1.20	1.14	0.02
Belgium	15	17	16	-1	1.07	1.02	1.09	0.02
Taiwan	14	13	17	-3	1.10	1.12	1.04	-0.06
Brazil	18	18	18	0	0.99	1.02	0.98	-0.01
Russia*	20	19	19	1	0.77	0.87	0.97	0.20
China	19	20	20	-1	0.86	0.84	0.78	-0.07
Mean:					1.17	1.18	1.19	
Standard Deviation:					0.21	0.20	0.17	

(* 1987 data are from West Germany and USSR respectively)

Table 5: Export Quality in 1987, 1997 and 2007.
Quality Rankings, BEC 3: Fuels and Lubricants

	Rank				Normalized Quality, World Average = 1			
country	1987	1997	2007	Change	1987	1997	2007	Change
USA	2	1	1	1	1.16	1.21	1.25	0.10
United Kingdom	8	5	2	6	1.05	1.11	1.16	0.11
Saudi Arabia	4	7	3	1	1.10	1.04	1.14	0.03
Netherlands	11	10	4	7	1.03	1.02	1.08	0.05
Australia	9	2	5	4	1.05	1.13	1.08	0.03
Oman	1	9	6	-5	1.35	1.02	1.06	-0.30
UAE	5	4	7	-2	1.08	1.12	1.04	-0.04
Algeria	3	8	8	-5	1.12	1.03	1.03	-0.09
Nigeria	20	6	9	11	0.84	1.08	1.01	0.17
Malaysia	10	3	10	0	1.04	1.13	1.00	-0.05
Iraq	7	16	11	-4	1.05	0.91	0.98	-0.07
Russia*	17	18	12	5	0.95	0.91	0.94	-0.01
Mexico	16	17	13	3	0.95	0.91	0.92	-0.03
Canada	18	12	14	4	0.93	0.99	0.92	-0.01
Venezuela	19	20	15	4	0.90	0.85	0.91	0.01
Norway	12	11	16	-4	1.02	1.01	0.90	-0.12
Iran	13	19	17	-4	1.01	0.88	0.90	-0.11
Qatar	15	14	18	-3	1.00	0.96	0.89	-0.10
Kuwait	6	15	19	-13	1.06	0.94	0.87	-0.19
Indonesia	14	13	20	-6	1.01	0.98	0.86	-0.15
Mean:					1.04	1.01	1.00	
Standard Deviation:					0.11	0.10	0.11	

(* 1987 data are from USSR)

Table 6: Export Quality in 1987, 1997 and 2007.
Quality Rankings, BEC 4: Capital Goods and Parts

	Rank				Normalized Quality, World Average = 1			
country	1987	1997	2007	Change	1987	1997	2007	Change
Switzerland	1	1	1	0	1.89	1.87	1.70	-0.19
Canada	2	4	2	0	1.40	1.38	1.41	0.01
Ireland	4	3	3	1	1.26	1.47	1.34	0.08
Sweden	7	2	4	3	1.22	1.49	1.32	0.11
Germany*	6	7	5	1	1.25	1.23	1.24	0.00
USA	5	6	6	-1	1.25	1.26	1.23	-0.02
United Kingdom	8	5	7	1	1.12	1.30	1.22	0.11
Netherlands	9	11	8	1	1.11	1.16	1.14	0.03
Belgium	10	8	9	1	1.06	1.23	1.14	0.07
France	3	10	10	-7	1.30	1.19	1.12	-0.19
Japan	11	9	11	0	0.95	1.21	1.06	0.11
Italy	12	13	12	0	0.94	1.00	1.02	0.08
Singapore	15	12	13	2	0.87	1.01	0.98	0.11
Mexico	13	15	14	-1	0.91	0.75	0.97	0.06
Malaysia	16	16	15	1	0.84	0.72	0.90	0.05
Rep. of Korea	18	14	16	2	0.55	0.80	0.88	0.33
Thailand	14	17	17	-3	0.90	0.72	0.81	-0.09
Taiwan	20	18	18	2	0.45	0.54	0.74	0.29
Hong Kong	17	19	19	-2	0.59	0.50	0.72	0.13
China	19	20	20	-1	0.49	0.36	0.52	0.03
Mean:					1.02	1.06	1.07	
Standard Deviation:					0.35	0.38	0.27	

(* 1987 data are from West Germany)

Table 7: Export Quality in 1987, 1997 and 2007.
Quality Rankings, BEC 5: Transport Equipment and Parts

	Rank				Normalized Quality, World Average = 1			
country	1987	1997	2007	Change	1987	1997	2007	Change
Austria	7	2	1	6	1.22	1.41	1.34	0.11
Sweden	1	1	2	-1	1.46	1.55	1.33	-0.12
United Kingdom	5	3	3	2	1.27	1.39	1.31	0.04
Spain	12	12	4	8	0.93	1.06	1.27	0.34
Canada	3	4	5	-2	1.28	1.30	1.26	-0.02
Hungary	19	15	6	13	0.64	0.92	1.25	0.60
Belgium	9	10	7	2	1.13	1.09	1.23	0.10
Germany*	2	7	8	-6	1.43	1.21	1.21	-0.22
USA	4	6	9	-5	1.27	1.23	1.20	-0.07
France	8	5	10	-2	1.22	1.24	1.19	-0.03
Italy	6	8	11	-5	1.23	1.19	1.18	-0.05
Netherlands	10	13	12	-2	1.08	1.02	1.12	0.03
Mexico	16	16	13	3	0.82	0.84	1.08	0.26
Japan	11	9	14	-3	0.97	1.12	1.07	0.11
Czech Rep.*	17	17	15	2	0.75	0.82	1.05	0.30
Brazil	13	11	16	-3	0.89	1.08	0.97	0.08
Poland	20	19	17	3	0.48	0.74	0.90	0.42
Rep. of Korea	14	14	18	-4	0.89	0.93	0.84	-0.05
Taiwan	15	18	19	-4	0.84	0.80	0.82	-0.01
China	18	20	20	-2	0.66	0.68	0.77	0.10
Mean:					1.02	1.08	1.12	
Standard Deviation:					0.28	0.24	0.18	

(* 1987 data are from West Germany and Czechoslovakia respectively)

Table 8: Export Quality in 1987, 1997 and 2007.
Quality Rankings, BEC 6: Consumer Goods

	Rank				Normalized Quality, World Average = 1			
country	1987	1997	2007	Change	1987	1997	2007	Change
Switzerland	1	1	1	0	2.07	2.48	2.72	0.65
Japan	6	2	2	4	1.48	1.88	1.99	0.51
United Kingdom	7	7	3	4	1.40	1.53	1.57	0.16
France	2	3	4	-2	1.80	1.74	1.56	-0.24
Italy	3	4	5	-2	1.58	1.68	1.51	-0.06
USA	8	10	6	2	1.35	1.34	1.41	0.07
Belgium	9	8	7	2	1.29	1.44	1.41	0.12
Ireland	5	5	8	-3	1.49	1.56	1.39	-0.10
Canada	12	12	9	3	1.15	1.25	1.35	0.20
Germany*	4	6	10	-6	1.54	1.54	1.35	-0.20
Netherlands	10	9	11	-1	1.25	1.38	1.30	0.05
Singapore	13	13	12	1	1.02	1.09	1.30	0.27
Spain	11	11	13	-2	1.24	1.31	1.19	-0.05
Hong Kong	17	18	14	3	0.87	0.84	1.09	0.22
Rep. of Korea	15	14	15	0	0.88	0.96	1.03	0.14
Mexico	19	16	16	3	0.76	0.90	1.02	0.26
Turkey	14	15	17	-3	0.94	0.94	0.89	-0.05
Taiwan	18	17	18	0	0.79	0.84	0.80	0.00
India	16	19	19	-3	0.88	0.79	0.75	-0.13
China	20	20	20	0	0.60	0.64	0.62	0.03
Mean:					1.22	1.31	1.31	
Standard Deviation:					0.38	0.45	0.46	

(* 1987 data are from West Germany)

relative quality has risen modestly in all other 1-digit BEC categories.

Some curious results in BEC 3: Fuels and Lubricants lead us to peer into the detailed calculations. Indonesia's recent low quality estimate is driven by the low relative quality of its coal and gas exports. Oman's high relative quality for 1987 comes from high quality estimates for its relatively modest exports of SITC codes 3341 ("Motor Spirit and Other Light Oils") and 3345 ("Lubricating Petroleum Oils and Other Heavy Oils"). The weight applied to Oman's quality estimates for these products depends not only on their importance in Oman's exports, but also on their importance in other countries' exports because we use Fisher Ideal price indexes. By 2007 Oman ceased to export SITC 3341 and its relative quality estimate in SITC 3345 had declined. Another interesting change in BEC 3 is the relative rise in U.S. export quality. Peering into the underlying data shows that this is driven by a rise in the relative U.S. export quality of SITC 3330 ("Petroleum Oils and Crude Oils"), which is generated by many countries reporting small volumes of imports from the U.S. with high unit values.

Our estimates call out for a comparison with the quality estimates of Hallak and Schott (2011) and Khandelwal (2010). We do this in Figure 9 using data from Table IV of Hallak and Schott and in Figure 10 using the median of HS 10-digit quality results for manufactured products generously provided by Amit Khandelwal. We take logs of our Table 2 results to make them more comparable with Hallak-Schott and demean all series.[25] Figure 9 compares our normalized quality estimates with Hallak–Schott in 1997 for the forty countries common to all three papers.[26] The correlation is extremely high, at 0.70, but there is a considerable difference in

[25] Khandelwal's quality estimates are not as directly comparable, since if translated to a CES framework they confound quality and the sensitivity of demand to price: see equation 15 of Khandelwal (2010).
[26] Hallak and Schott's quality estimates for each country are linear trends, so it is a simple matter to back out the implied 1997 results.

Figure 9: Comparison With Hallak and Schott (2011)

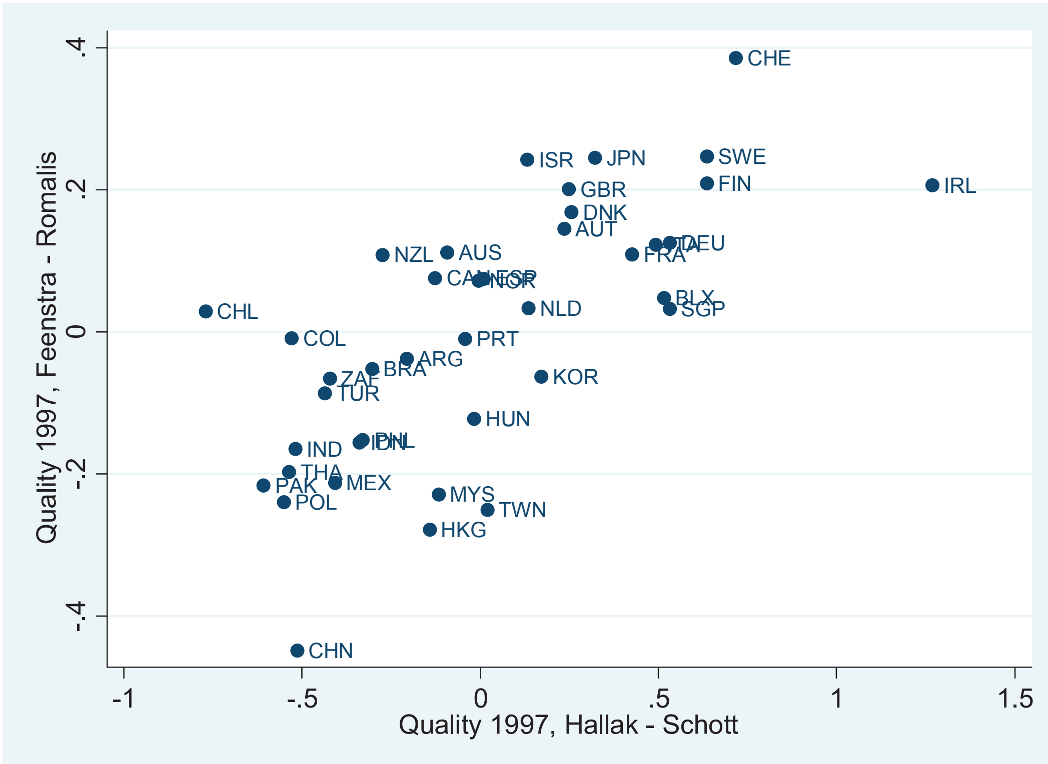
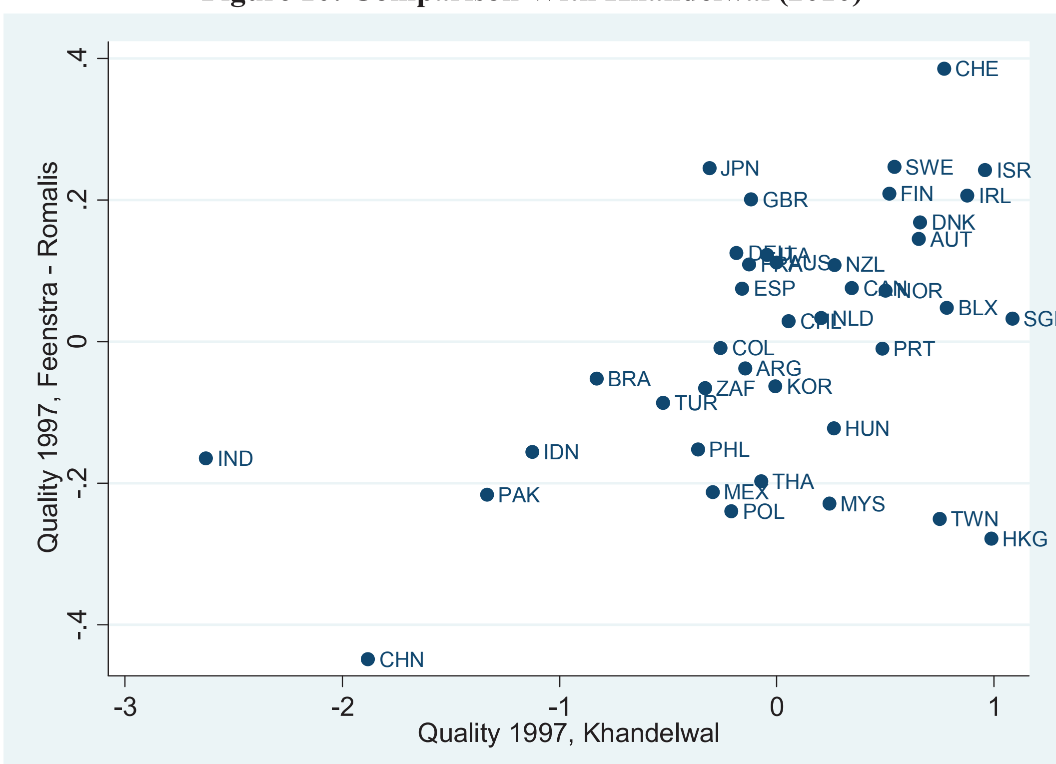


Figure 10: Comparison With Khandelwal (2010)



the dispersion of the two sets of estimates. The standard deviation of the Hallak-Schott quality estimates is 0.45, compared with 0.18 for our matching estimates. The lower dispersion of our estimates likely reflects our use of world-wide trade data in all products rather than just U.S. manufacturing imports, but may also be partly due to different estimation and aggregation procedures.

Figure 10 provides the equivalent comparison with Khandelwal (2010). The correlation between the two sets of estimates is high, at 0.49, but the higher dispersion of Khandelwal's estimates (the standard deviation is 0.77) cannot be directly compared with the other estimates.[27] There is one interesting feature of this comparison though relating to population. A close look at Figure 10 suggests that while the quality estimates of our paper are closely related to income per capita, Khandelwal's are closely related to population – the countries to the right of the figure have fairly small populations while those to the left have large populations. This association is driven by the use of population as a proxy for variety.[28] Less obviously, the Hallak-Schott estimates are closely related to the manufacturing trade balance, which is a key component of their measure of demand. These associations are made crystal-clear in Table 9, which reports regressions of the three sets of export quality estimates plus our import-quality and terms of trade estimates on three country-level variables: log per capita income from the PWT; log population; and the manufacturing trade balance from the UN's Comtrade database divided by manufacturing value added from the World Bank's World Development Indicators.[29]

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[27] See note 25.

[28] Following Khandelwal (2010), we have used the estimated labor force in each SITC industry and country as a proxy for export variety, as explained in note 12. While this proxy enters into the gravity equation (19), and thereby affect the estimated parameters from this equation, it does not otherwise enter into the formulas for quality or quality-adjusted prices.

[29] Since Hallak and Schott report trend values of quality, we take an average of the manufacturing trade balance to value added ratio over their 1989 to 2003 sample period.

Table 9: Comparison of Quality Estimates for 1997

	Hallak and Schott (2011)	Khandelwal (2010)	Feenstra and Romalis (this paper)		
Dependent variable:	*Export quality*	*Export quality*	*Export quality*	*Import quality*	*Terms of trade*
Independent variables:					
Log GDP Per Capita	0.32	0.31	0.14	0.01	-0.07
	(0.05)	(0.06)	(0.03)	(0.02)	(0.02)
Log Population	-0.08	-0.37	-0.01	-0.01	-0.03
	(0.03)	(0.04)	(0.02)	(0.01)	(0.01)
Manufacturing Trade Balance / Value Added	0.84	0.17	0.09	0.03	-0.04
	(0.08)	(0.11)	(0.06)	(0.04)	(0.03)
Observations	38	38	38	38	38
R-squared	0.88	0.92	0.54	0.15	0.43

Notes: Standard errors are reported in parentheses. The ratio of the manufacturing trade balance to manufacturing value added variable has been averaged over Hallak and Schott's (2011) 1989 to 2003 sample period. We lose two countries, Israel and Taiwan, due to missing manufacturing value added data in the World Development Indicators.

All three export quality estimates are strongly positively correlated with per capita income. Khandelwal's estimates exhibit a very strong relationship to country population, while Hallak and Schott's estimates are moderately correlated with population and our estimates are uncorrelated with population. The Hallak-Schott quality estimates are very strongly correlated with the manufacturing trade balance, while Khandelwal's and our export quality estimates are only slightly correlated with that balance. Our import quality estimates are not significantly correlated with any of the three variables.[30] Finally, our quality-adjusted terms of trade estimates for these countries are negatively correlated with per capita income and population, but are not associated with the manufacturing trade balance.

[30] The relationship between import quality and log GDP per capita evident in the middle panel of Figure 7 is not present in 1997, but a regression over the full sample of about 200 countries in 2007 generates a significant positive coefficient on log GDP per capita of 0.017 (s.e. 0.004). Interestingly, the full sample of countries in 1987 yields a marginally significant *negative* coefficient on log GDP per capita: -0.010 (0.006).

8. Conclusions

Our goal has been to adjust observed trade unit values for quality so as to estimate quality-adjusted prices in trade. We achieve this goal by explicitly modeling the quality choice by exporting firms in an environment where consumers have non-homothetic tastes for quality. We find a greater preference for quality in richer countries, consistent with Hallak (2006). Our key parameter estimate of the elasticity of quality with respect to the quantity of inputs almost always lies between zero and unity, as required by our model. This implies that only a fraction of observed import unit-value differences are due to quality, with the remainder reflecting differences in quality-adjusted import prices.

Our estimates of the elasticity of substitution between different varieties of the same SITC 4-digit products are substantially higher than in Broda and Weinstein (2006). As a result, the observed differences in export unit-values are attributed predominantly to quality, with very small remaining difference in quality-adjusted export prices. The quality-adjusted terms of trade therefore declines with country income, at least in 2007, reflecting rich countries' preferences for higher quality and therefore higher quality-adjusted prices. In that year variation in the quality-adjusted terms of trade is only one-half as large as that in the unadjusted ratio of export to import unit-value indexes.

In contrast to existing literature, which has tended to focus on exports and often for one destination market (e.g. the United States), we construct quality and quality-adjusted prices for imports and exports (at the BEC 1-digit level) for all countries included in the Penn World Table (PWT). These estimates will be used to construct an output-based measure of real GDP, reflecting countries terms of trade, in the "next generation" of the PWT.

Appendix A: Gravity Equation

Combining (16) and (18), we readily solve for:

$$\overline{P_i^k} = \left(\overline{p_i^k} \Big/ \left(\kappa_1^k \, \overline{p_i^{*k}} \right)^{\alpha^k \theta} \right)^{\frac{1}{1+\alpha^k \theta(\sigma-1)}} \left(\frac{Y^k P^{k(\sigma-1)}}{\kappa_2^k tar_i^k f_i^k} \right)^{\frac{\alpha^k \theta}{1+\alpha^k \theta(\sigma-1)}} . \tag{A1}$$

We obtain the gravity equation by solving for the CES price index in (17). To this end, we first replace the quality-adjusted prices appearing in (17) by using their solution in (A1). Second, that solution from (A1) is substituted into (18) to obtain exports,

$$\left(\frac{X_i^k / \kappa_2^k f_i^k}{M_i(\varphi_i / w_i)^{\gamma}} \right) = \left(\frac{\overline{p_i^k}}{\left(\kappa_1^k \, \overline{p_i^{*k}} \right)^{\alpha^k \theta}} \right)^{-\frac{(\sigma-1)(1+\gamma)}{[1+\alpha^k \theta(\sigma-1)]}} \left(\frac{Y^k P^{k(\sigma-1)}}{\kappa_2^k tar_i^k f_i^k} \right)^{\frac{(1+\gamma)}{[1+\alpha^k \theta(\sigma-1)]}} . \tag{A2}$$

Third, we solve for the export probabilities $(\hat{\varphi}_i^k / \varphi_i)^{-\gamma}$ appearing in (17) using (14),

$$\left(\frac{\hat{\varphi}_i^k}{\varphi_i} \right)^{-\gamma} = \left[\frac{\overline{X_i^k} / \kappa_2^k tar_i^k f_i^k}{M_i(\varphi_i / w_i)^{\gamma}} \right]^{\frac{-\gamma}{(1+\gamma)}} \left(\frac{w_i}{\varphi_i} \right)^{-\gamma} . \tag{A3}$$

We now follow the same steps as in Chaney (2008), which means that we substitute (A2) into (A3) to obtain an expression for the export probabilities that depends on the c.i.f. prices, f.o.b. prices, trade costs, income and the price index P^k itself. That solution is substituted back into (17) to solve for the CES price index in terms of those other variables. That solution is:

$$Y^k P^{k(\sigma-1)} = \kappa_2^k \left(\frac{Y^k}{\kappa_2^k \widetilde{M}^k} \right)^{\frac{[1+\alpha^k \theta(\sigma-1)]}{(1+\gamma)}} , \tag{A4}$$

where κ_2^k is defined by (14), and $\widetilde{M}^k$ is defined by,

$$\widetilde{M}^k \equiv \sum_i M_i \left(\frac{\varphi_i}{w_i}\right)^{\gamma} \left(\overline{p_i^k} \Big/ \left(\kappa_1^k \overline{p_i^{*k}} \right)^{\alpha^k \theta} \right)^{\frac{-(\sigma-1)(1+\gamma)}{[1+\alpha^k \theta(\sigma-1)]}} \left(tar_i^k f_i^k \right)^{-\left[\frac{\gamma-\alpha^k\theta(\sigma-1)}{1+\alpha^k\theta(\sigma-1)}\right]}. \quad \text{(A5)}$$

The gravity equation is obtained by substituting (A5) back into (A2):

$$\left(\frac{X_i^k}{M_i(\varphi_i / w_i)^{\gamma}} \right) = \left(\frac{\overline{p_i^k}}{\left(\kappa_1^k \overline{p_i^{*k}} \right)^{\alpha^k \theta}} \right)^{\frac{-(\sigma-1)(1+\gamma)}{[1+\alpha^k \theta(\sigma-1)]}} \left(tar_i^k f_i^k \right)^{-\left[\frac{\gamma-\alpha^k\theta(\sigma-1)}{1+\alpha^k\theta(\sigma-1)}\right]} \left(\frac{Y^k}{\widetilde{M}^k} \right). \qquad \text{(A6)}$$

Then using the specification for fixed costs in (19), we obtain (20)–(21), where the term $(Y^k / uv^k)^{\beta_0}$ from (19) cancels out. The exponents in the gravity equation appear complex, but in fact, are not too different from those in Chaney (2008) as can be seen by allowing $\alpha^k \theta \to 1$. In this limit we have $\kappa_1^k \to (\sigma - 1) / \sigma$, which can be ignored as a constant. Then the price term in (A6) approaches the ratio $(\overline{p_{it}^k} / \overline{p_{it}^{*k}})$, which from (8) equals the iceberg trade costs τ_i^k. The exponent of that term approaches $-(\sigma - 1)(1 + \gamma) / \sigma$. In contrast, Chaney (2008) finds that the exponent of iceberg trade costs is simply the Pareto parameter $-\gamma$. This difference between our gravity equation and Chaney's is explained by the fact that we have allowed the fixed costs of exporting to depend on the productivity of the firm. The second term on the right of (A6) is the fixed costs of exporting, inclusive of one plus the *ad valorem* tariff, raised to a power that is again similar to that in Chaney (2008) when $\alpha^k \theta \to 1$. So difference between Chaney's gravity equation and what we find when $\alpha^k \theta \to 1$ is due to our modeling of fixed costs as depending on productivity and the adjustment to this term due to *ad valorem* tariffs.

Appendix B: Data

(i) Trade Data: We obtain all bilateral international trade values and quantities for the SITC Revision 2 classification from the United Nation's COMTRADE database. Where possible, quantities for a given SITC code are converted into common units. Where this is not possible, each combination of SITC code and unit of quantity is treated as a separate product.

(ii) Distance Data: The distance between countries is measured as the great-circle distance between the capital cities of those two countries.

***(iii) Tariff Data*:** We obtain tariff schedules from five primary sources: (i) raw tariff schedules from the TRAINS database accessed via the World Bank's WITS website date back as far as 1988 for some countries; (ii) manually entered tariff schedules published by the International Customs Tariffs Bureau (BITD) dating back as far as the 1950's;[31] (iii) U.S. tariff schedules from the U.S. International Trade Commission from 1989 onwards;[32] (iv) derived from detailed U.S. tariff revenue and trade data from 1974 to 1988 maintained by the Center for International Data at UC Davis; and (v) the texts of preferential trade agreements primarily sourced from the WTO's website, the World Bank's Global Preferential Trade Agreements Database, or the Tuck Center for International Business Trade Agreements Database. For the U.S., specific tariffs have been converted into ad-valorem tariffs by dividing by the average unit value of matching imported products. Due to the difficulties of extracting specific tariff information for other countries and matching it to appropriate unit values, only the ad-valorem component of their tariffs are used. The overwhelming majority of tariffs are ad-valorem. Switzerland is a key exception here, with tariffs being specific. We proxy Swiss tariffs with tariffs of another EFTA

[31] Most tariff schedules can be fairly readily matched to the SITC classification.
[32] See Feenstra, Robert C., John Romalis, and Peter K. Schott (2002) for a description of U.S. tariff data for 1989 onwards.

member (Norway). We aggregate MFN and each non-MFN tariff program[33] to the 4-digit SITC Revision 2 level by taking the simple average of tariff lines within each SITC code.

Tariff schedules are often not available in each year, especially for smaller countries. Updated schedules are more likely to be available after significant tariff changes. Rather than replacing "missing" MFN tariffs by linearly interpolating observations, missing observations are set equal to the nearest preceding observation. If there is no preceding observation, missing MFN tariffs are set equal to the nearest observation.

Missing non-MFN tariff data (other than punitive tariffs applied in a handful of bilateral relationships) are more difficult to construct for two reasons: (i) it is often not published in a given tariff schedule; (ii) preferential trade agreements have often been phased in. To address this we researched the text of over 100 regional trade agreements and Generalized System of Preferences (GSP) programs to ascertain the start date of each agreement or program and how the typical tariff preference was phased in. To simplify our construction of missing preferential tariffs we express observed preferential tariffs as a fraction of the applicable MFN tariff. We fill in missing values of this fraction based on information on how the tariff preferences were phased in. Preferential tariffs are then constructed as the product of this fraction and the MFN tariff. We then keep the most favorable potentially applicable preferential tariff. Punitive non-MFN tariff levels tend not to change over time (though the countries they apply to do change). We replace missing observations in the same way we replace missing MFN tariff observations.

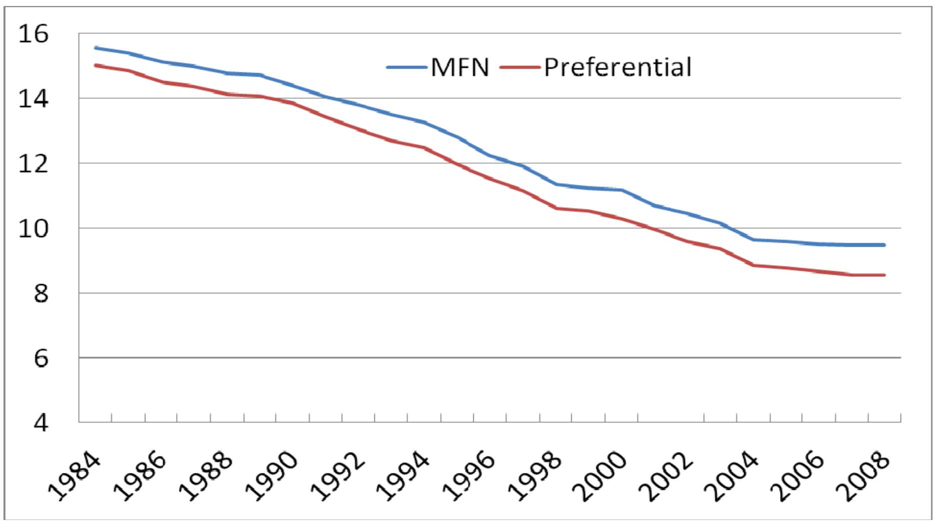
The evolution of a simple average of these MFN and most favorable preferential tariffs from 1984 to 2008 is summarized in Appendix Figure 1. Since MFN tariffs apply to most

[33] Multiple preferential tariffs may be applicable for trade in a particular product between two countries. Since the most favorable one may change over time, we keep track of each potentially applicable tariff program.

bilateral relationships, the average "Preferential" tariff is only slightly lower than the average MFN tariff.

(iv) Quality-Adjusted Unit Values: The quality estimates shown by (27) and (28) depend on the c.i.f. and f.o.b. unit values, but two-thirds of the bilateral, 4-digit SITC trade flows in our

Appendix Figure 1: Typical MFN and Preference-adjusted Tariff*



*Notes: Simple average across all potential bilateral trade relationships and products. If no tariff preference applies the MFN tariff is used.

Comtrade data that have quantity information are missing one unit-value or the other. So while in our estimation we use only the observations where both the c.i.f. and f.o.b. unit values are available, to construct the quality-adjusted prices we want to fill in for the missing c.i.f. data. To achieve this we use the structure of our model, where from (8) the ratio of c.i.f. to f.o.b. prices is proportional to the iceberg trade costs. Replacing the c.i.f. and f.o.b. prices by their unit values as in (23), we use estimates from the preliminary regression:

$$\ln(uv_{igt}^{k} / uv_{igt}^{*k}) = \beta_{0g} + \beta_{1g} \ln dist_{i}^{k} + u_{igt}^{k} - u_{igt}^{*k},$$

to form an estimate of the predicted iceberg trade costs, $\exp(\hat{\beta}_{0g} + \hat{\beta}_{1g} \ln dist_{i}^{k})$. Then when an f.o.b. unit value uv_{it}^{*k} is available but not the c.i.f. unit value, we can impute the c.i.f. unit value by $uv_{it}^{*k} \times \exp(\hat{\beta}_{0g} + \hat{\beta}_{1g} \ln dist_i^k)$, and when only the c.i.f. unit value is available then we impute the f.o.b. unit value by $uv_{it}^{k} / \exp(\hat{\beta}_{0g} + \hat{\beta}_{1g} \ln dist_i^k)$.

(v) Language Data: Data on 6,909 spoken languages in almost all countries is published in M. Paul Lewis (2009) and available online at www.ethnologue.com (Ethnologue). We collected data on the number of speakers in each country of languages that are spoken by 0.5 percent or more of the local population, and on immigrant languages that are either spoken by more than 0.1 percent of the local population or are an official language. Official language data is primarily collected from the Central Intelligence Agency's "The World Factbook" (2012), supplemented by data from Lewis (2009) when The World Factbook does not list official languages. Spoken and official languages are then classified by Matthew S. Dryer and Martin Haspelmath (2011) *The World Atlas of Language Structures* (WALS) into languages, and progressively broader groupings: language genus, language sub-family, and language family. For most languages language sub-family is not defined, so we collect data on language, language genus and language family. For example, Swedish belongs to the Germanic language genus and the Indo-European language family. In this way we capture the fact that Swedish is closer to German than it is to French, which belongs to the Romance genus, and closer to French than to Swahili which belongs to the Niger-Congo language family. This process is rendered difficult for three reasons. Firstly, Ethnologue is more liberal at classifying a dialect as a separate language than is WALS, so we have to look to Ethnologue's more detailed but less systematic classification scheme to infer what WALS language Ethnologue is referring to. Secondly, Ethnologue and WALS sometimes use different names for the same languages, which have to be reconciled by searching their lists of alternative names. Finally, WALS is incomplete, so we infer a WALS classification using Lewis's classification.

Appendix C: Estimation

For convenience, we omit the goods subscript g in what follows, though all parameters and equations differ by SITC good.

To utilize the GMM methodology introduced by Feenstra (1994), we need to develop the supply side in more detail. The c.i.f. and f.o.b. prices shown in (8) depend on the iceberg and the specific transport costs. The former depends on one plus the *ad valorem* tariffs, denoted by tar_{it}^{k}, and we model both costs as also depending on the distance from country *i* to *k* and the aggregate physical export quantity $(X_{it}^{k} / uv_{it}^{k})$, in a log-linear fashion:

$$\ln \tau_{it}^{k} = \eta_t + \eta_0 \ln tar_{it}^{k} + \eta_1 \ln dist_i^{k} + \eta_2 \ln(X_{it}^{k} / uv_{it}^{k}) + \xi_{1it}^{k}, \qquad \text{(C1)}$$

$$\ln T_{it}^{k} = \chi_t + \chi_1 dist_i^{k} + \chi_2 \ln(X_{it}^{k} / uv_{it}^{k}) + \xi_{2it}^{k}. \qquad \text{(C2)}$$

We are including the quantity exported $(X_{it}^{k} / uv_{it}^{k})$ to reflect possible congestion (or scale economies) in shipping, and also so that our model nests that used in Feenstra (1994). We treat the random errors ξ_{1it}^{k} and ξ_{2it}^{k} as independent of ε_{it}^{k}.

Notice from (8a) and (8b) that $\overline{p_{it}^{k}} \Big/ \overline{p_{it}^{*k}} = \kappa_3^{k} \tau_{it}^{k}$, $\kappa_3^{k} \equiv \left[\left(\frac{1}{1-\alpha^{k}\theta}\right)\left(\frac{\sigma}{\sigma-1}\right)\right] \Big/ \left[\left(\frac{1}{1-\alpha^{k}\theta}\right)\left(\frac{\sigma}{\sigma-1}\right)-1\right]$. Combining this with $\overline{p_{it}^{k}} = \tau_{it}^{k}(\overline{p_{it}^{*k}} + T_{it}^{k})$, (C1), (C2) and (23), we write an inverse supply curve using a similar linear combination of c.i.f. and f.o.b. unit values that appear in the demand equation (25):

$$\begin{aligned}\left[\left(\ln uv_{it}^{k} - \ln uv_{jt}^{k}\right) - \theta\left(\ln uv_{it}^{*k} - \ln uv_{jt}^{*k}\right)\right] &= (\eta_0 - 1)(\ln tar_{it}^{k} - \ln tar_{jt}^{k}) \\ + \omega_1 (dist_i^{k} - dist_j^{k}) + \omega_2 \left[\ln(X_{it}^{k} / uv_{it}^{k}) - \ln(X_{jt}^{k} / uv_{jt}^{k})\right] &+ (\tilde{\xi}_{it}^{k} - \tilde{\xi}_{jt}^{k}),\end{aligned} \qquad \text{(C3)}$$

where $\omega_i \equiv (\eta_i - \theta\chi_i)$, i=1,2, and $\tilde{\xi}_{it}^{k} \equiv (\xi_{1it}^{k} + u_{it}^{k}) - \theta(\xi_{2it}^{k} - u_{it}^{*k})$ incorporates the measurement error in (23). We rewrite (C3) slightly by shifting the export unit values and values to the left:

$$\left[(1+\omega_2)\left(\ln uv_{it}^{k}-\ln uv_{jt}^{k}\right)-\theta\left(\ln uv_{it}^{*k}-\ln uv_{jt}^{*k}\right)\right]-\omega_2(\ln X_{it}^{k}-\ln X_{it}^{k})$$
$$=(\eta_0-1)(\ln tar_{it}^{k}-\ln tar_{jt}^{k})+\omega_1\,(dist_i^{k}-dist_j^{k})+(\tilde{\xi}_{it}^{k}-\tilde{\xi}_{jt}^{k}). \tag{C4}$$

We combine this supply curve with the gravity equation (25), rewritten slightly as:

$$\ln X_{it}^{k}-\ln X_{jt}^{k}+A^{k}\left[\left(\ln uv_{it}^{k}-\ln uv_{jt}^{k}\right)-\alpha^{k}\theta\left(\ln uv_{it}^{*k}-\ln uv_{jt}^{*k}\right)\right]=\delta_0(\ln L_{it}-\ln L_{jt})$$
$$+\delta_i-\delta_j-\sum_{n=1}^{N}B^{k}\beta_n(F_{ni}^{k}-F_{nj}^{k})-C^{k}\left(\ln tar_{it}^{k}-\ln tar_{jt}^{k}\right)+\tilde{\varepsilon}_{it}^{k}-\tilde{\varepsilon}_{jt}^{k}, \tag{C5}$$

where $\tilde{\varepsilon}_{it}^{k}\equiv(\varepsilon_{it}^{k}+A^{k}u_{it}^{k}-A^{k}\alpha^{k}\theta u_{it}^{*k})$ includes the measurement error in (23), and A^{k}, B^{k}, C^{k} are given by (27) using $\alpha^{k}\equiv 1+\lambda\ln\left(\sum_{t=1}^{T}\frac{1}{T}\frac{RGDPL_t^{k}}{RGDPL_t^{US}}\right)$ as the time-average of (9b).

Taking the product of (C4) and (C5) and dividing by $A^{k}(1+\omega_2)$, we obtain:

$$\left(\ln uv_{it}^{k}-\ln uv_{jt}^{k}\right)^2$$
$$=\left[\alpha^{k}\theta+\frac{\theta}{(1+\omega_2)}\right]\left(\ln uv_{it}^{k}-\ln uv_{jt}^{k}\right)\left(\ln uv_{it}^{*k}-\ln uv_{jt}^{*k}\right)-\frac{\alpha^{k}\theta^2}{(1+\omega_2)}\left(\ln uv_{it}^{*k}-\ln uv_{jt}^{*k}\right)^2$$
$$+\frac{\omega_2}{A^{k}(1+\omega_2)}\left(\ln X_{it}^{k}-\ln X_{jt}^{k}\right)^2+\left(\frac{\omega_2}{(1+\omega_2)}-\frac{1}{A^{k}}\right)\left(\ln X_{it}^{k}-\ln X_{jt}^{k}\right)\left(\ln uv_{it}^{k}-\ln uv_{jt}^{k}\right)$$
$$+\left(\frac{\theta}{A^{k}(1+\omega_2)}-\frac{\alpha^{k}\omega_2\theta}{(1+\omega_2)}\right)\left(\ln X_{it}^{k}-\ln X_{jt}^{k}\right)\left(\ln uv_{it}^{*k}-\ln uv_{jt}^{*k}\right)+Controls_{it}^{k}+\mu_{it}^{k}, \tag{C6}$$

with the control terms,

$$Controls_{it}^{k}\equiv\frac{1}{A^{k}(1+\omega_2)}\left[(\eta_0-1)(\ln tar_{it}^{k}-\ln tar_{jt}^{k})+\omega_1(dist_i^{k}-dist_j^{k})\right]\times$$
$$\left[\delta_0(\ln L_{it}-\ln L_{jt})+\delta_i-\delta_j-\sum_{n=1}^{N}B^{k}\beta_n(F_{ni}^{k}-F_{nj}^{k})-C^{k}\left(\ln tar_{it}^{k}-\ln tar_{jt}^{k}\right)\right] \tag{C7}$$

and the error term,

$$\mu_{it}^{k}=\frac{(\tilde{\xi}_{it}^{k}-\tilde{\xi}_{jt}^{k})}{A^{k}(1+\omega_2)}\left[\delta_i-\delta_j+\delta(\ln L_{it}-\ln L_{jt})-\sum_{n=1}^{N}B^{k}\beta_n(F_{ni}^{k}-F_{nj}^{k})+\tilde{\varepsilon}_{it}^{k}-\tilde{\varepsilon}_{jt}^{k}\right]$$
$$+\frac{(\tilde{\varepsilon}_{it}^{k}-\tilde{\varepsilon}_{jt}^{k})}{A^{k}(1+\omega_2)}\left[(\eta_0-1)(\ln tar_{it}^{k}-\ln tar_{jt}^{k})+\omega_1\,(dist_i^{k}-dist_j^{k})\right].$$

We treat the country fixed effects, sectoral labor force, distance, tariffs and language variables for the fixed costs of exporting as exogenous, so they are uncorrelated with the demand and supply shocks. We further assume that the supply and gravity shocks are uncorrelated in (C4) and (C5), so that $E\mu_{it}^k = 0$ for each source country *i* and destination *k*. This is the moment condition that we use to estimate (C6). This equation is simplified using $\zeta^k = \gamma / [\alpha^k \theta(\sigma - 1)]$ and so $\zeta^k \alpha^k = \gamma / [\theta(\sigma - 1)] = \zeta^{US} \alpha^{US} = \zeta^{US}$, since $\alpha^{US} \equiv 1$ by normalization. It follows that $\gamma = \zeta^{US} \theta(\sigma - 1)$, and then from (C5) we obtain $A^k = (\sigma - 1)[1 + \zeta^{US} \theta(\sigma - 1)] / [1 + \alpha^k \theta(\sigma - 1)]$ and $B^k = [\zeta^{US} \theta(\sigma - 1) - \alpha^k \theta(\sigma - 1)] / [1 + \alpha^k \theta(\sigma - 1)]$. Substituting these relations into (C6), we obtain an equation that is nonlinear in the parameters θ and σ.

For estimation, we average the variables in (C6) over time, which eliminates the time subscript and gives a cross-country regression that can be estimated with nonlinear least squares (NLS). A final challenge is to incorporate the source country fixed effects $(\delta_i - \delta_j)$ interacted with distance and tariffs as appear in (C7). The list of countries varies by product, so it is difficult to incorporate these interactions directly into the NLS estimation. Instead, we first regress *all other* variables in (C6) on the source country fixed effects and their interaction terms, and then estimate (C6) using the residuals obtained from these preliminary regressions. The source country fixed effects are needed to control for the measurement errors in the c.i.f. and f.o.b. unit values, shown in (23), which we assume are independent of each other and of the export values. Then the variance of the measurement errors appears in the error term after averaging over time, and the source country fixed effects absorb these variances.

Because the GMM estimation is performed after eliminating the source country fixed effects and their interactions, we do not obtain the coefficients of those terms. Likewise, we do not recover the estimates of the other control terms in (C7). So a second-stage estimation is performed to obtain these coefficients. In particular we substitute the estimates $\hat{\sigma}, \hat{\theta}, \hat{\gamma}$ and $\hat{\lambda}$ into (26) to obtain the coefficients $\hat{A}^k$, $\hat{B}^k$. Then writing $\ln[M_{it}(w_{it} / \phi_{it})^{-\gamma}] = \delta_i + \delta_0 \ln L_{it} + \varepsilon_{it}^k$ more generally as $\ln[M_{it}(w_{it} / \phi_{it})^{-\gamma}] = \delta_{it} + \varepsilon_{it}^k$, and also allowing the coefficients β_{0t} and β_{nt} to vary over time, the gravity equation (25) is run again over time for each SITC good:

$$\ln(X_{it}^k / Y_t^k) + \hat{A}^k \left[\left(\ln \overline{uv_{it}^k} - \ln \overline{uv_{jt}^k} \right) - \hat{\alpha}^k \hat{\theta} \left(\ln \overline{uv_{git}^{*k}} - \ln \overline{uv_{gjt}^{*k}} \right) \right]$$
$$= \delta_{it} + \delta_t^k - \hat{B}^k \sum_{n=1}^{4} \beta_{nt} F_{ni}^k - \left\{ \beta_{5t} \left[\frac{\hat{\sigma}(1+\hat{\gamma})}{1+\hat{\alpha}^k \hat{\theta}(\hat{\sigma}-1)} \right] - 1 \right\} \ln tar_{it}^k + \varepsilon_{it}^k, \tag{C8}$$

where the error term $\hat{\varepsilon}_{it}^k$ includes $\tilde{\varepsilon}_{it}^k$ plus the sampling error in the coefficients $\hat{\sigma}, \hat{\theta}, \hat{\gamma}$ and $\hat{\lambda}$, and $\delta_t^k \equiv -\ln M_t^k$. Running (C8) as a fixed-effects regression for each SITC good and each year, we obtain $\hat{\delta}_{it}$ and $\hat{\delta}_t^k$, so that $\hat{M}_t^k = e^{-\hat{\delta}_t^k}$ is the estimate of market potential. In addition, we estimate the coefficients $\beta_{nt}, n = 1,...,5,$ from which we construct the fixed costs of exporting $\exp\left(\sum_{n=1}^{4} \hat{\beta}_{nt} F_{ni}^k\right)$ that is used when obtaining the quality-adjusted prices. It turns out that multicollinearity in the language variables leads to some SITC industries where the estimates of $\beta_{nt}, n = 1,...,5,$ are quite large positive or negative values, leading to extreme values of the estimated fixed costs. Accordingly, we omitted 5% of the bilateral fixed cost estimates (i.e. the top and bottom 2.5 percentiles) and replaced these with average values.

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