International Prices and Endogenous Quality*

Robert C. Feenstra,
University of California, Davis
and NBER

John Romalis
University of Chicago, Booth School of Business
and NBER

March 2012

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 employ a demand system to model consumer demand in which quality and quantity multiply each other in the utility function. In that case, the quality choice by firms is the solution to a cost-minimization sub-problem. We estimate the gravity equation for this system using detailed bilateral trade data for over 150 countries during 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.

* The authors thank George Dan, John Lennox, Anson Soderbery and Greg Wright for excellent research assistance, as well as Juan Carlos Hallak, Robert Inklaar, Marcel Timmer and seminar participants at the NBER for helpful comments. Financial support from the National Science Foundation is gratefully acknowledged.

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, which 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 with that goal 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

[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 that assumption which is 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 (2008), Crozet, Head and Mayer (2009), Johnson (2009) and Mandel (2009).

procedure 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 the early work by Rodriguez (1979), we allow quality to multiply quantity in the utility function, leading to a sub-problem of quality choice for the firm: to minimize the average cost of 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 the aggregate input price, as well as the specific transport costs to the destination market. 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 input price.

In order to implement this measure of quality, we therefore need accurate information on the 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 input prices 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 that arises in the endogenous-quality model and compare it to Chaney (2008). In sections 4 and 5, we estimate the gravity equation using detailed bilateral trade data at the 4-digit SITC digit level (nearly 1,000 products per year) for over 150 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, and due to the fact that we model both the intensive and extensive margins of trade. 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, but the quality differentials are much lower than the raw differences in unit values. Countries' quality-adjusted terms of trade, though varying widely, exhibit little relationship to their level of development, in contrast to their unadjusted terms of trade. 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 their impact on real GDP.

2. Optimal Quality Choice

Consumer Problem

Suppose that consumers in country *k* have available $i=1,\ldots,N^k$ 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 will suppose that the demand for the products in country *k* arises from a sub-utility function $U(z_1^k q_1^k,\ldots,z_{N^k}^k q_{N^k}^k)$ for each sector, where quality z_i^k multiplies the quantity q_i^k. Later we will specialize to the CES form:

$$U(z_1^k q_1^k,\ldots,z_{N^k}^k q_{N^k}^k) = \sum_{i=1}^{N^k} \left(z_i^k q_i^k\right)^{(\sigma-1)/\sigma}, \qquad \sigma > 1. \tag{1}$$

This sub-utility function for each sector can be nested in an overall utility function that is non-homothetic, allowing for income difference across countries to affect sectoral demand, as in Fieler (2011) for example.[3] We do not make this relationship explicit, however, but focus on a single sector in the theory while applying our estimation to many such sectors.

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

We suppose there are both specific and *ad valorem* trade costs between the countries. Specific trade costs are given by T_i^k, which can depend on the distance to the destination market *k*. One plus the *ad valorem* trade costs are denoted by τ_i^k, which can also depend on distance and on the *ad valorem* tariffs. For convenience we sometimes refer to τ_i^k as just tariffs, but they also include insurance and all other trade costs proportional to value. We assume that the *ad valorem* charges are applied to the value *inclusive* of the specific trade costs.[4] Then letting p_i^{*k} denote the exporters' f.o.b. price, the tariff-inclusive c.i.f. price is $p_i^k \equiv \tau_i^k (p_i^{*k} + T_i^k)$.

Thus, consumers in country *k* are presented with a set of $i = 1,\ldots,N^k$ varieties, with characteristics z_i^k and prices p_i^k, and then choose the optimal quantity of each variety. It will be convenient to work with the *quality-adjusted, tariff-inclusive c.i.f. prices*, which are defined by the upper-case $P_i^k \equiv p_i^k / z_i^k = \tau_i^k (p_i^{*k} + T_i^k) / z_i^k$. The higher is overall product quality z_i^k, *ceteris paribus*, the lower are the quality-adjusted prices P_i^k. The consumer maximizes utility subject to the budget constraint $\sum_{i=1}^N \tau_i^k (p_i^{*k} + T_i^k) q_i^k \le Y^k$. The Lagrangian for country *k* is,

$$
\begin{aligned}
L &= U(z_1^k q_1^k, \ldots, z_{N^k}^k q_{N^k}^k) + \lambda[Y^k - \textstyle\sum_{i=1}^{N^k} \tau_i^k (p_i^{*k} + T_i^k) q_i^k] \\
&= U(Q_1^k, \ldots, Q_{N^k}^k) + \lambda(Y^k - \textstyle\sum_{i=1}^{N^k} P_i^k Q_i^k), \qquad (2)
\end{aligned}
$$

where the second line of (2) follows by defining $Q_i^k \equiv z_i^k q_i^k$ as the *quality-adjusted demand*, and also using the quality-adjusted prices $P_i^k \equiv \tau_i^k (p_i^{*k} + T_i^k) / z_i^k$. This change of variables in the

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[4] Most jurisdictions 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 accessible at http://export.customsinfo.com/ and http://export.gov/logistics/eg_main_018142.asp.

Lagrangian makes it clear that instead of choosing q_i^k given c.i.f. prices $\tau_i^k(p_i^{*k} + T_i^k)$ and quality z_i^k, we can instead think of the representative consumer as choosing Q_i^k given quality-adjusted c.i.f. prices P_i^k, i =1,…,N^k. Let us denote the solution to problem (2) by $Q_i(P^k, Y^k)$, i = 1,…,N^k, where P^k is the vector of quality-adjusted prices.

Firms' Problem

We now add the subscript j for firms, while i denotes their country of origin, so that *(i,j)* denotes a unique variety. We will denote the range of firms exporting from country i to k by $j = 1,..., N_i^k$. Firms make the optimal choice of the quality z_{ij}^k to send to country k. We assume that the input x_{ij}^k needed to produce one unit of a good with product quality z_{ij}^k arises from a Cobb-Douglas function:

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

where $0 < \theta < 1$ reflects diminishing returns to quality and φ_{ij} denotes the productivity of firm j in country i. We think of x_{ij}^k as an aggregate of inputs, such as high and low-skilled labor, capital and entrepreneurial ability, as in Verhoogen (2008), and denote its aggregate price by 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. The profits from exporting to country k are:

$$\max_{p_{ij}^{*k}, z_{ij}^k} [p_{ij}^{*k} - c_{ij}(z_{ij}^k, w_i)]q_{ijt}^k = \max_{p_{ij}^{*k}, z_{ij}^k} \left[\frac{p_{ij}^{*k}}{z_{ij}^k} - \frac{c_{ij}(z_{ij}^k, w_i)}{z_{ij}^k} \right] Q_{ij}^k(P^k, Y^k)$$
$$= \max_{P_{ij}^k, z_{ij}^k} \left\{ \frac{P_{ij}^k}{\tau_i^k} - \frac{[c_{ij}(z_{ij}^k, w_i) + T_i^k]}{z_{ij}^k} \right\} Q_{ij}^k(P^k, Y^k) \qquad (5)$$

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 $P_{ij}^k \equiv \tau_i^k (p_{ij}^{*k} + T_i^k) / z_{ij}^k$ along with demands Q_{ij}^k. The latter transformation relies on our assumption that *prices and characteristics are chosen simultaneously,* as well as our assumption that quality multiplies quantity in the utility function, but does not rely on the CES form in (1).

It is immediate that to maximize profits in (5), the firms must choose z_{ij}^k to minimize $[c_{ij}(z_{ij}^k, w_i) + T_i^k] / z_{ij}^k$, which is interpreted as *minimizing the average variable cost per unit of quality, inclusive of specific trade costs*. The same optimality condition appears in Rodriguez (1979) who also assumes that quantity multiplies quality in the utility function. Differentiating this objective with respect to z_{ij}^k, we obtain the first-order condition:

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

so that the average variable cost per unit of quality equals the marginal cost when average variable costs are minimized. The second-order condition for this cost-minimization problem is that $\partial^2 c_{ij} / \partial(z_{ij}^k)^2 > 0,$ so there must be increasing marginal costs of improving quality. An increase in the distance to the destination market raises T_i^k, so to satisfy (6) firms will choose a higher quality z_{ij}^k, as readily shown from $\partial^2 c_{ij} / \partial(z_{ij}^k)^2 > 0.$ This is the well-known "Washington apples" effect, whereby higher quality goods are sent to more distant markets.

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 then be written as:

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

Conveniently, the Cobb-Douglas production function and specific trade costs give us a log-linear form for the optimal quality choice. This relationship is suggestive of the log-linear regressions between exporter f.o.b. prices and the distance to destination markets, which typically give a *positive* coefficient on distance (Hummels and Skiba, 2004) that is interpreted as quality and taken as evidence of the "Washington applies" effect. We see from (7) that more distant markets, with higher transport costs T_i^k, will indeed have higher quality, but that log quality is only a fraction $\theta < 1$ of the log transport costs. In addition, higher firm productivity φ_{ij} leads to lower effective wages (w_i / φ_{ij}) and also leads to higher quality. Finally, substituting (7) into the cost function (4), we immediately obtain $c_{ij}(z_{ij}^k, w_i) = [\theta / (1-\theta)]T_i^k$. Thus, the marginal costs of production are proportional to the specific trade costs, as we shall use below.

Now suppose that the CES utility function in (1) applies. Solving (5) for the optimal choice of the f.o.b. price, we obtain the familiar markup,

$$(p_{ij}^{*k} + T_i^k) = [c_{ij}(z_{ij}^k, w_i) + T_i^k]\left(\frac{\sigma}{\sigma - 1}\right). \tag{8}$$

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 $c_{ij}(z_{ij}^k, w_i) = [\theta / (1-\theta)]T_i^k$, we readily solve for the f.o.b. and tariff-inclusive c.i.f. prices as:

$$p_{ij}^{*k} = T_i^k \left[\left(\frac{1}{1-\theta} \right) \left(\frac{\sigma}{\sigma - 1} \right) - 1 \right] \equiv \overline{p_i^{*k}}, \tag{9a}$$

$$p_{ij}^{k} = \tau_i^k T_i^k \left[\left(\frac{1}{1-\theta} \right) \left(\frac{\sigma}{\sigma - 1} \right) \right] \equiv \overline{p_i^{k}}. \tag{9b}$$

Thus, both the f.o.b. and c.i.f. prices vary across destination markets *k* in direct proportion to the specific transport costs to each market, but are independent of the productivity of the firm *j*, as indicated by the notation $\ln \overline{p_i^{*k}}$ and $\ln \overline{p_i^{k}}$. This result is obtained because more efficient firms sell higher quality goods, leading to constant prices in each destination market.[5]

Combining (7) and (9a) we obtain:

$$z_{ij}^k = \left[\overline{p_i^{*k}} \Big/ (w_i / \varphi_{ij}) \right]^{\theta} \kappa_1, \tag{10}$$

where κ_1 is a parameter depending on θ and σ.[6] Thus, quality z_{ij}^k depends on the ratio of the f.o.b. price $\overline{p_i^{*k}}$ to the productivity-adjusted input price (w_i / φ_{ij}) of the exporter, with an exponent that depends on the ease with which quality is varied in the production function (3). It follows that the quality-adjusted price $P_{ij}^k = \overline{p_i^k} / z_{ij}^k$ is:

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

Since from (9) 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 utility function in (1), sales

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[5] 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.

[6] $\kappa_1 = \{\theta(\sigma - 1) / [1 + \theta(\sigma - 1)]\}^{\theta}$.

depend on the quality-adjusted price with elasticity $(1-\sigma)$, and from (11), the price depends on productivity with elasticity $-\theta$, so that firms' sales depends on productivity with elasticity $\theta(\sigma-1)$. 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 $\zeta = \gamma / [\theta(\sigma-1)]$. We can see that in order for our model to mimic the Melitz-Chaney model, we need to have $\theta \to 1$, so that prices would decline at the rate $\gamma / (\sigma-1)$ as in Chaney (2008). Setting $\theta = 1$ is not permitted because the quality approaches infinity in (7), but we will occasionally let $\theta \to 1$ 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 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 input price 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:

$$\hat{P}_i^k = \left(\overline{p_i^k} \Big/ \overline{p_i^{*k}}^{\theta} \right) \left(\frac{w_i}{\hat{\varphi}_i^k} \right)^{\theta} / \kappa_1 . \qquad (11')$$

We let $\hat{Q}_i^k$ denote the quantity of exports for this marginal firm so that $\hat{X}_i^k / \tau_i^k = \hat{P}_i^k \hat{Q}_i^k / \tau_i^k$ is the export revenue net of the *ad valorem* trade costs. It is immediate from the CES markup rule in (8) that profits earned by the firm in (5) equal $(\hat{X}_i^k / \sigma\tau_i^k)$, which must cover fixed costs. We shall assume that a firm with the productivity-adjusted input price (w_i / φ_{ij}) has fixed costs of $(w_i / \varphi_{ij}) f^k$ to sell to destination market *k*, where f^k is common to all firms exporting to *k*. That is, the firm's productivity applies equally well to variable and fixed costs.[7] This assumption is also made by Bilbiie, Ghironi, and Melitz (2007), though in their case, productivity is equal across firms. Then the zero-cutoff-profit (ZCP) condition to determine productivity is:[8]

$$\frac{\hat{X}_i^k}{\sigma\tau_i^k} = \frac{w_i f^k}{\hat{\varphi}_i^k} \quad . \tag{12}$$

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 a common dispersion parameter across countries. The Pareto distribution is $G_i(\varphi) = 1 - (\varphi / \varphi_i)^{-\gamma}$, where the location parameter $\varphi_i \le \varphi$ 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

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[7] If we alternatively modeled the fixed costs as just $w^k f^k$, then the cutoff firms from all countries *i* facing the same *ad valorem* trade costs to a destination country *k* would have identical quality-adjusted prices, so that any observed differences in their unit-values would be *entirely* explained by quality. To avoid this case we must assume that at least *some part* of the fixed costs depend on firms' productivity.

[8] Because we are treating the *ad valorem* trade costs τ_i^k as including tariffs, we divide revenue by σ and τ_i^k on the left of (12) to obtain profits. This differs from the usual treatment of iceberg trade costs, where it is assumed that the shipping costs accrue to the exporter so that revenue is only divided by σ to obtain profits. This slight difference in the ZCP condition in (12) will influence the form that τ_i^k enters the gravity equation: see note 18.

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:

$$X_i^k \equiv M_i \int_{\hat{\varphi}_i^k}^{\infty} X_{ij}^k g_i(\varphi)d\varphi = M_i \int_{\hat{\varphi}_i^k}^{\infty} \hat{X}_i^k (P_{ij}^k / \hat{P}_i^k)^{1-\sigma} g_i(\varphi)d\varphi$$

$$= M_i \hat{X}_i^k \left[\int_{\hat{\varphi}_i^k}^{\infty} \left(\frac{\hat{\varphi}_i^k}{\varphi} \right)^{\theta(1-\sigma)} g_i(\varphi)d\varphi \right] = M_i \hat{X}_i^k \left[\frac{\gamma}{\gamma - \theta(\sigma-1)} \right] \left(\frac{\hat{\varphi}_i^k}{\varphi_i} \right)^{-\gamma},$$

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

$$\left(\frac{w_i}{\hat{\varphi}_i^k} \right)^{1+\gamma} = \frac{\overline{X_i^k} / \tau_i^k f^k}{M_i(\varphi_i / w_i)^{\gamma}} \left[\frac{\gamma - \theta(\sigma-1)}{\gamma\sigma} \right]. \qquad (13)$$

Substituting this solution for productivity-adjusted wages into (11'), we readily obtain an 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 country *k* is:

$$\overline{P_i^k} \equiv \left[\int_{\hat{\varphi}_i^k}^{\infty} P_{ij}^k(\varphi)^{(1-\sigma)} \frac{g_i(\varphi)}{\left[1-G_i(\hat{\varphi}_i^k)\right]} d\varphi \right]^{\frac{1}{1-\sigma}} = \left[\frac{\gamma}{\gamma-\theta(\sigma-1)} \right]^{\frac{1}{1-\sigma}} \hat{P}_i^k, \qquad (14)$$

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[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.

as obtained by substituting for $P_{ij}^{k}(\varphi_{ij})$ from (11) and computing the integral for $\gamma > \theta(\sigma - 1)$. 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'), (13) and (14), we therefore obtain,[10]

$$\overline{P_i^k} = \left(\overline{p_i^k} \Big/ \overline{p_i^{*k}}^{\theta} \right) \left(\frac{X_i^k / \tau_i^k f^k}{M_i (\varphi_i / w_i)^{\gamma}} \right)^{\frac{\theta}{(1+\gamma)}} \kappa_2 . \qquad (15)$$

Notably, an increase in exports to a market, given the mass of firms, *raises* the relative quality-adjusted price. That is because an increase in relative exports means that less-efficient firms must be exporting to that market, and therefore average quality is falling. That relationship sounds contrary to the demand-side intuition discussed in section 1: given nominal prices, high relative sales to a market should mean high relative quality. In fact, that intuition continues to hold in our model, and we shall use it below in conjunction with (15) to solve for the quality-adjusted prices and to develop our estimating equations.

Import Demand

Returning to the zero-profit cutoff condition in (12), while the firm-level sales $\hat{X}_i^k$ are not observed in our data, they equal CES demand from the utility function in (1). In particular, the ratio of firm-level sales from country *i* to market *k* is simply $\hat{X}_i^k = (\hat{P}_i^k / P^k)^{-(\sigma-1)} Y^k$, where P^k is the exact price index for the CES utility function:

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

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[10] $\kappa_2 = \{[\gamma - \theta(\sigma-1)] / \sigma\gamma]\}^{\theta/(1+\gamma)} / \kappa_1$.

using (14). Also using aggregate exports along with (13), demand is re-expressed as:

$$\left(\frac{X_i^k / \tau_i^k f^k}{M_i(\varphi_i / w_i)^\gamma}\right) = \left(\frac{\overline{P_i^k}}{P^k}\right)^{-(\sigma-1)(1+\gamma)} \left(\frac{Y^k}{\tau_i^k f^k}\right)^{(1+\gamma)} \left[\frac{\gamma - \theta(\sigma-1)}{\sigma\gamma}\right]^\gamma . \qquad (16)$$

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.[11] The presence of this term plagues 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 (19).[12] 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 ultimately rely on a similar assumption, but first show how the mass of exporters can be eliminated by combining the demand-side relation (16) with the supply-side relation (15) to solve for the quality-adjusted price.

Combining (15) and (16), we readily solve for: [13]

$$\overline{P_i^k} = \left(\overline{p_i^k} \Big/ \overline{p_i^{*k}}^{\theta}\right)^{\frac{1}{1+\theta(\sigma-1)}} \left(\frac{Y^k P^{k(\sigma-1)}}{\tau_i^k f^k}\right)^{\frac{\theta}{1+\theta(\sigma-1)}} \kappa_3 . \qquad (17)$$

Notice that the mass of exporters M_i no longer enters this expression, but the quality-adjusted

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[11] The mass of firms is multiplied by minimum productivity relative to wages, $(\varphi_i / w_i)^\gamma$. With equilibrium wages reflecting productivity, we could expect this term to be close to unity.

[12] Given that either greater variety or higher quality have similar effects on demand, one might try to construct price indexes that correct for both. We do not do that here – focusing on only quality – because while a variety correction is quite well-understood on the import side it is more controversial on the export side (see Feenstra, 2010), and more importantly, because national statistical agencies do not attempt to correct for variety.

[13] $\kappa_3 = \{[\gamma - \theta(\sigma-1)] / \sigma\gamma\}^{\theta/[1+\theta(\sigma-1)]} / \kappa_1^{1/[1+\theta(\sigma-1)]}$.

price still depends on the price index P^k in the destination country. Following Chaney (2008), we can solve for this price index (see Appendix A) and use that solution in (17) to obtain:

$$\overline{P_i^k} = \left(\frac{\overline{p_i^k}}{\overline{p_i^{*k}}^{\theta}} \right)^{\frac{1}{[1+\theta(\sigma-1)]}} \left(\tau_i^k f^k\right)^{\frac{-\theta}{[1+\theta(\sigma-1)]}} \left(\frac{Y^k}{M^k}\right)^{\frac{\theta}{(1+\gamma)}}, \tag{18}$$

where M^k denotes the “supplier access” or “market potential” for country *k*,[14]

$$M^k \equiv \left\{ \sum_i M_i \left(\frac{\varphi_i}{w_i}\right)^{\gamma} \left(\frac{\overline{p_i^k}}{\overline{p_i^{*k}}^{\theta}}\right)^{\frac{-(\sigma-1)(1+\gamma)}{[1+\theta(\sigma-1)]}} \left(\tau_i^k f^k\right)^{-\left[\frac{\gamma-\theta(\sigma-1)}{1+\theta(\sigma-1)}\right]} \right\}. \tag{19}$$

So rather than the quality-adjusted price depending on the mass of firms M_i exporting from country *i*, we instead end up in (19) with the *worldwide* mass of firms M^k exporting to country *k.* In the summation in (19), the mass of firms M_i exporting from each country is weighted by terms that inversely reflect the trade costs between *i* and *k* and the fixed costs of exporting.[15]

Taking the ratio of (18) for two exporters *i* and *j* selling to the same country *k*, we obtain:

$$\frac{\overline{P_i^k}}{\overline{P_j^k}} = \left(\frac{\overline{p_i^k} \big/ \tau_i^k \, \overline{p_i^{*k}}^{\theta}}{\overline{p_j^k} \big/ \tau_j^k \, \overline{p_j^{*k}}^{\theta}} \right)^{\frac{1}{[1+\theta(\sigma-1)]}} = \left(\frac{\overline{p_i^k}}{\overline{p_j^k}}\right)^{\rho}, \quad \rho \equiv \frac{(1-\theta)}{[1+\theta(\sigma-1)]}, \tag{20}$$

where the second equality follows by substituting for the *ad valorem* trade costs τ_i^k from (9), which are proportional to the ratio of the c.i.f. to the f.o.b. prices. Thus, we find that that all unknown terms like fixed costs f^k and market potential M^k no longer enter the ratio of quality-

[14] “Supplier access” is the term used by Redding and Venables (2004) in a monopolistic competition model with homogeneous firms. It can also be thought of as “market potential” from the buyer’s point of view.

[15] Using (9), the price ratio $\overline{p_i^k} \big/ \overline{p_i^{*k}}^{\theta}$ is proportional to trade costs $\tau_i^k T_i^{k(1-\theta)}$, so the presence of this price ratio in (19) also reflects inverse trade costs.

adjusted prices, which becomes a simple root of the ratio of c.i.f. prices in (20), with $0 < \rho < 1$. This formulation makes it clear that the disparity in nominal c.i.f. prices across source countries – such as found by Schott (2004) – is diminished in the quality-adjusted prices. Stated differently, the ratio of quality itself is also a simple root of the ratio of c.i.f. prices:

$$\frac{\overline{z_i^k}}{\overline{z_j^k}} \equiv \left(\frac{\overline{p_i^k} \big/ \overline{P_i^k}}{\overline{p_j^k} \big/ \overline{P_j^k}} \right) = \left(\frac{\overline{p_i^k}}{\overline{p_j^k}} \right)^{1-\rho} .$$

Thus, from the combination of the supply-side relation (15) and demand-side relation (16) we obtain a very sharp solution for quality and the quality-adjusted prices for two countries selling to the same destination market. Indeed, we will argue that (20) is all that is needed to measure the quality-adjusted *export* prices of countries, which has been the focus in the literature. To obtain the quality-adjusted *import* prices we will need to take into account the additional terms in (18), however, including market potential.

In order to estimate the model parameters we set (15) equal to (18) to eliminate the quality-adjusted price, and also now add a time subscript *t*, obtaining the gravity equation:

$$\left(\frac{X_{it}^k}{M_{it}(\varphi_{it} / w_{it})^{\gamma}} \right) = \left(\frac{\overline{p_{it}^k}}{\overline{p_{it}^{*k}}^{\theta}} \right)^{\frac{-(\sigma-1)(1+\gamma)}{[1+\theta(\sigma-1)]}} \left(\tau_{it}^k f_t^k \right)^{-\left[\frac{\gamma-\theta(\sigma-1)}{1+\theta(\sigma-1)} \right]} \left(\frac{Y_t^k}{M_t^k} \right). \qquad (21)$$

This formulation of the gravity equation has as its first term the quality-adjusted price from (11), without any adjustment for wages. The exponent of this term combines the elasticity of substitution σ and the Pareto parameter γ because we have solved for both the intensive and extensive margins of trade.[16] The second term is the fixed costs of exporting raised to a power

[16] It can be shown that if we had not made the fixed costs of exporting depend on productivity in our model, then as $\theta \to 1$ the exponent of this term would approach $-\gamma$, as Chaney (2008) finds applies to trade costs.

similar to that the gravity equation of Chaney (2008), but here the fixed costs are also multiplied by the *ad valorem* costs of trade.[17] The third term is also very similar to that in Chaney, whom we have followed in deriving the "market potential."[18]

We can get even closer to Chaney's form of the gravity equation by substituting for the *ad valorem* trade costs, which from (9) are proportional to the ratio of c.i.f. and f.o.b. prices, $\tau_i^k = \kappa_4 \overline{p_i^k} / \overline{p_i^{*k}}$.[19] Substituting above, we obtain the gravity equation:

$$\left(\frac{X_{it}^k}{M_{it}(\varphi_{it}/w_{it})^{\gamma}}\right) = \underbrace{\left(\frac{\overline{p_{it}^k}}{\overline{p_{it}^{*k}}}\right)^{-\gamma}}_{\propto(\tau_i^k)^{-\gamma}} \left(\kappa_4 f_t^k\right)^{-\left[\frac{\gamma-\theta(\sigma-1)}{1+\theta(\sigma-1)}\right]} \left(\frac{Y_t^k}{M_t^k}\right) \underbrace{\left(\overline{p_{it}^k}\right)^{\frac{-(1-\theta)(\sigma-1)(1+\gamma)}{[1+\theta(\sigma-1)]}}}_{\propto \overline{P_i^k}^{-(\sigma-1)(1+\gamma)}} \tag{22}$$

In the first term in (22) we obtain the *ad valorem* trade costs $\tau_i^k \propto \overline{p_i^k} / \overline{p_i^{*k}}$ raised to the power $-\gamma$, which is identical to Chaney. The second term again are again fixed costs raised to a power similar to that in Chaney, and third term is also similar. What is new in (22) is the last term appearing on the right, which is the c.i.f. price raised to a power. Notice that as $\theta \to 1$, this term vanishes and we have essentially the same gravity equation as Chaney. For $\theta < 1$ we interpret this term by taking ratio between two countries *i* and *j* selling to the same destination *k*. In that case, since from (20):

$$\left(\overline{p_{it}^k} / \overline{p_{jt}^k}\right)^{\frac{-(1-\theta)(\sigma-1)(1+\gamma)}{[1+\theta(\sigma-1)]}} = \left(\overline{P_{it}^k} / \overline{P_{jt}^k}\right)^{-(\sigma-1)(1+\gamma)},$$

---

[17] Chaney finds that the fixed costs are raised to the power $-[\gamma-(\sigma-1)]/(\sigma-1)<0$, which is similar to the second term in (22) when $\theta \to 1$. The *ad valorem* costs of trade multiply the fixed costs because we divide revenue by the former in the ZCP condition (12), so it is readily seen that $\tau_{it}^k f^k$ can be combined on the right of (12).

[18] We can obtain the term like what Chaney (2008) uses for market potential by setting $w_i = \varphi_i$ in (19), re-writing the prices appearing in (19) as in note 16, and replacing M_i with L_i, the population of the exporting country.

[19] $\kappa_4 = \left[\left(\frac{1}{1-\theta}\right)\left(\frac{\sigma}{\sigma-1}\right)-1\right] \Big/ \left[\left(\frac{1}{1-\theta}\right)\left(\frac{\sigma}{\sigma-1}\right)\right]$.

we interpret the c.i.f. prices raised to the power $-\frac{(1-\theta)(\sigma-1)(1+\gamma)}{[1+\theta(\sigma-1)]}$ as proportional to the quality-adjusted prices to the power $-(\sigma-1)(1+\gamma)$, which is exactly the same way that the quality-adjusted prices enter the demand equation (16). In other words, we interpret the final term in (22) as once again reflecting the demand-side intuition that higher exports, on the left, imply a lower quality-adjusted price on the right, *ceteris paribus*.

To estimate (22), notice that the mass of potential exporters now re-enters on the left, so we will model this term as depending on country fixed effects and the estimated labor force L_{it} involved in the production of exports in each sector and country:[20]

$$\ln[M_{it}(w_{it}/\phi_{it})^{-\gamma}] = \alpha_i + \alpha \ln L_{it} + \varepsilon_{it}^k, \tag{23}$$

where ε_{it}^k is a random error. Substituting this equation into (22), taking the difference between two countries *i* and *j* selling to destination *k*, we obtain:

$$\begin{aligned}\ln X_{it}^k - \ln X_{jt}^k = -(\gamma + A)\left(\ln \overline{p_{it}^k} - \ln \overline{p_{jt}^k}\right) + \gamma\left(\ln \overline{p_{it}^{*k}} - \ln \overline{p_{jt}^{*k}}\right) \\ + \alpha_i - \alpha_j + \alpha(\ln L_{it} - \ln L_{jt}) + \varepsilon_{it}^k - \varepsilon_{jt}^k,\end{aligned} \tag{24}$$

where, $$A \equiv \frac{(1-\theta)(\sigma-1)(1+\gamma)}{[1+\theta(\sigma-1)]} = \frac{(1-\theta)}{\theta}\frac{\gamma(1+\gamma)}{(\zeta+\gamma)}. \tag{25}$$

It is evident that in this gravity equation, the tariff-inclusive c.i.f. prices $\overline{p_{it}^k}$ enter with a negative coefficient $-(\gamma + A)$ that exceeds in absolute value the positive coefficient γ on the f.o.b. price. This pattern reflect the appearance of the quality-adjusted price as the last term in (21). Both the c.i.f. and f.o.b. prices are endogenous and subject to measurement error, so we use the GMM

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[20] 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.

methodology of Feenstra (1994) to estimate (24). But obtaining unbiased estimates of γ and A does not allow us to identify all the parameters entering the expression for A. So we rely on the estimates of $\zeta = \gamma / [\theta(\sigma - 1)]$ that Chaney (2008) obtains from regressions of firm-rank on firm-size within SITC sectors to rewrite A as shown in second equality in (25).[21] Given these values for ζ, along with our estimates of γ and θ, we can use the estimated value of A to obtain the elasticity of substitution σ. We will discuss our estimation technique in more detail in the next section, after first reviewing the data.

4. Data and Estimation

Data

The primary dataset used is the United Nations' Comtrade database. We obtain bilateral f.o.b. prices of traded goods by calculating the unit value of each bilateral trade observation at the 4-digit SITC industry level, as reported by the exporting country. By focusing on the exporters' reports we ensure that these values are calculated prior to the inclusion of any costs of shipping the product. The bilateral c.i.f. prices are calculated similarly using importers' trade reports. Since these values include the costs of shipping, we need only to add the value of any tariff on the good to produce a tariff-inclusive c.i.f. price. To do this we obtain tariff values 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.

To our knowledge, we are the first to make detailed use of both the importing country's

[21] We thank Thomas Chaney for providing these estimates for 3-digit SITC Rev. 3 sectors, which we concorded to 3-digit SITC Rev. 2 sectors. His pooled estimate of ζ over all sectors is 2.4, which is considerably higher than estimated by, for example, Di Giovanni *et al* (2011).

reported c.i.f. prices and the exporting country's reported f.o.b. prices when estimating the demand and quality of traded goods. The independent variation in these data is essential to identifying the negative coefficient on the c.i.f. price and the positive coefficient on the f.o.b. price in (24). But must be admitted that there is a large amount of measurement error in prices – actually unit values – from the Comtrade database. This measurement error is apparent by merging the trade values and quantities reported by exporters and importers (by year and 4-digit SITC), and then finding that only about *one-third* of the resulting observations have both the c.i.f. and f.o.b. unit values for that trade flow; the remaining two-thirds of disaggregate trade flows are only reported by one partner or the other. Therefore, the estimation of our system is done over this one-third of observations that have both unit values.

Furthermore, when both the c.i.f. and f.o.b. data are available, 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). 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_{it}^{k} and uv_{it}^{*k}, are related to the true c.i.f. and f.o.b. prices by:

$$\ln uv_{it}^{k} = \ln \overline{p_{it}^{k}} + u_{it}^{k} \quad \text{and} \quad \ln uv_{it}^{*k} = \ln \overline{p_{it}^{*k}} + u_{it}^{*k}, \tag{26}$$

where u_{it}^{k} and u_{it}^{*k} are the measurement errors that are independent of each other and have variances σ^{k} and σ_{i}^{*}, 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 it is crucial that the errors are independent for this claim to hold, which is therefore an identifying assumption.

Estimation

Our goal is to estimate equation (24) to obtain estimates of θ, σ and θ, while recognizing that the expenditure shares and prices appearing there are endogenous. To control for this endogeneity we will modify the GMM methodology introduced by Feenstra (1994), which exploits the moment condition that the error in demand and supply are uncorrelated. That assumption could be violated when quality is present, however, since a change in quality could act as shift to both supply and demand. While that criticism can be made of Feenstra (1994) and Broda and Weinstein (2003), it does not apply here because we have explicitly modeled quality choice. To complete our model, we need to develop the supply side in more detail.

The c.i.f. and f.o.b. prices shown in (9) depend on the *ad valorem* 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 trade quantity $\overline{Q_{it}^{k}} \equiv X_{it}^{k} / \overline{P_{it}^{k}}$ exported in a log-linear fashion:

$$\ln \tau_{it}^{k} = \beta_t + \beta_0 \ln tar_{it}^{k} + \beta_1 \ln dist_i^{k} + \beta_2 \ln \overline{Q_{it}^{k}} + \delta_{1it}^{k}, \qquad (27)$$

$$\ln T_{it}^{k} = \eta_t + \eta_1 dist_i^{k} + \eta_2 \ln \overline{Q_{it}^{k}} + \delta_{2it}^{k}. \qquad (28)$$

We are including the quantity exported $\overline{Q_{it}^{k}}$ to reflect possible congestion (or scale economies) in shipping, and also so that our model nests that used in Feenstra (1994). We suppose that both transport costs depend on global trends β_t and η_t, reflecting productivity, and treat the random errors δ_{1it}^{k} and δ_{2it}^{k} as independent of ε_{it}^{k}.

Combining (27)-(28) with (9), we write an inverse supply curve using the same linear combination of c.i.f. and f.o.b. prices that appear in the demand equation (24):

$$\begin{aligned}(\gamma + A)\left(\ln \overline{p_{it}^{k}} - \ln \overline{p_{jt}^{k}}\right) - \gamma\left(\ln \overline{p_{it}^{*k}} - \ln \overline{p_{jt}^{*k}}\right) = \omega_0(\ln tar_{it}^{k} - \ln tar_{it}^{k}) \\ +\omega_1(dist_i^k - dist_j^k) + \omega_2(\ln \overline{Q_{it}^{k}} - \ln \overline{Q_{jt}^{k}}) + (\delta_{it}^{k} - \delta_{jt}^{k}),\end{aligned} \tag{28}$$

where $\omega_0 \equiv (\gamma + A)\beta_0$, $\omega_i \equiv [(\gamma + A)\beta_i + A\eta_i]$, i=1,2, and $\delta_{it}^{k} \equiv [(\gamma + A)\delta_{1it}^{k} + A\delta_{2it}^{k}]$. Equations (24) and (28) are similar to the system in Feenstra (1994), except for three features: (i) both the c.i.f. and f.o.b. prices appear; (ii) the presence of tariffs, labor and distance as control variables; (iii) we do not express the system in first-differences over time, because we want to retain distance as a variable that is important for the choice of quality. We rewrite (28) by using the export value to replace the quantity $\overline{Q_{it}^{k}} \equiv X_{it}^{k} / \overline{P_{it}^{k}}$, and using (20) to replace $\overline{P_{it}^{k}}$, obtaining:

$$\begin{aligned}(\gamma + A + \omega_2\rho)\left(\ln \overline{p_{it}^{k}} - \ln \overline{p_{jt}^{k}}\right) - \gamma\left(\ln \overline{p_{it}^{*k}} - \ln \overline{p_{jt}^{*k}}\right) = \omega_0(\ln tar_{it}^{k} - \ln tar_{jt}^{k}) \\ +\omega_1(dist_i^k - dist_j^k) + \omega_2(\ln X_{it}^{k} - \ln X_{it}^{k}) + (\delta_{it}^{k} - \delta_{jt}^{k}).\end{aligned} \tag{29}$$

Multiplying (24) and (29) gives a lengthy equation, reported in the Appendix C, which has an error depending on the product $(\varepsilon_{it}^{k}\delta_{it}^{k})$ and variables that are the second moments and cross-moments of the data. This is the analogue to the demand and supply system in Feenstra (1994), extended here to endogenous quality choice. Feenstra (1994) assumed that the supply shocks are uncorrelated with the demand shocks. That assumption is unlikely to hold with unobserved quality, however, since a change in quality could shift both supply and demand. But in this paper, the errors ε_{it}^{k} and δ_{it}^{k} are the residuals in demand (24) and supply (28) after taking into account quality. The assumption that ε_{it}^{k} and δ_{it}^{k} are uncorrelated therefore seems much more acceptable, and is the basis for the GMM estimation.

5. Parameter Estimates

Estimation is performed for each 4-digit SITC Revision 2 goods (which we also refer to as industries) 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 844 industries as shown in the first row of Table 1.[22] The median estimate of σ is 6.94, with an inadmissible value less than unity in only one case; the median estimate of γ is 8.40, with an inadmissible negative value in only three cases; and the median estimate of θ is 0.58, with an inadmissible value less than zero or greater than unity in only five cases, though some of these inadmissible cases refer to the same SITC industry as reported in Table 1.[23] Excluding these inadmissible cases, the range of estimates for the parameters are illustrated in Figures 1-3.

When there are two or more units of measurement used in an industry, we have estimated the system separately over the various units since otherwise the unit values are non-comparable. In the second row of Table 1 we report the median estimates over 716 SITC industries where we use only the primary unit of measurement (nearly always kilograms). The median estimates do not change by very much, so in the ensuing work we retain the secondary units of measurement, which treating the primary and secondary units as different SITC goods.

Our median estimate for σ is higher than the median estimates obtained by Broda and Weinstein (2006). Based on Table IV of their paper, we would expect to obtain a median elasticity at the SITC 4-digit level between 2.5 and 2.8, and we instead get 6.9. There are two possible explanations for this difference. The first is that we have estimated the system over all

[22] We drop SITC industries where the NSL estimates for γ or A fail to converge.

[23] For inadmissible or non-convergent parameter values, we replace the parameter estimates with the median estimate from the same 3-digit or 2-digit SITC industry.

Table 1: Median Parameter Estimates

GMM Estimation Method with:	Number of SITC industries	σ	γ	θ
Dropping SITC with < 50 observations	844	6.94	8.40	0.58
Number of inadmissible parameters	6	1	3	5
+ Dropping SITC with secondary unit of measurement	716	7.12	8.79	0.58
Number of inadmissible parameters	1	0	0	1

Figure 1: Estimates of σ

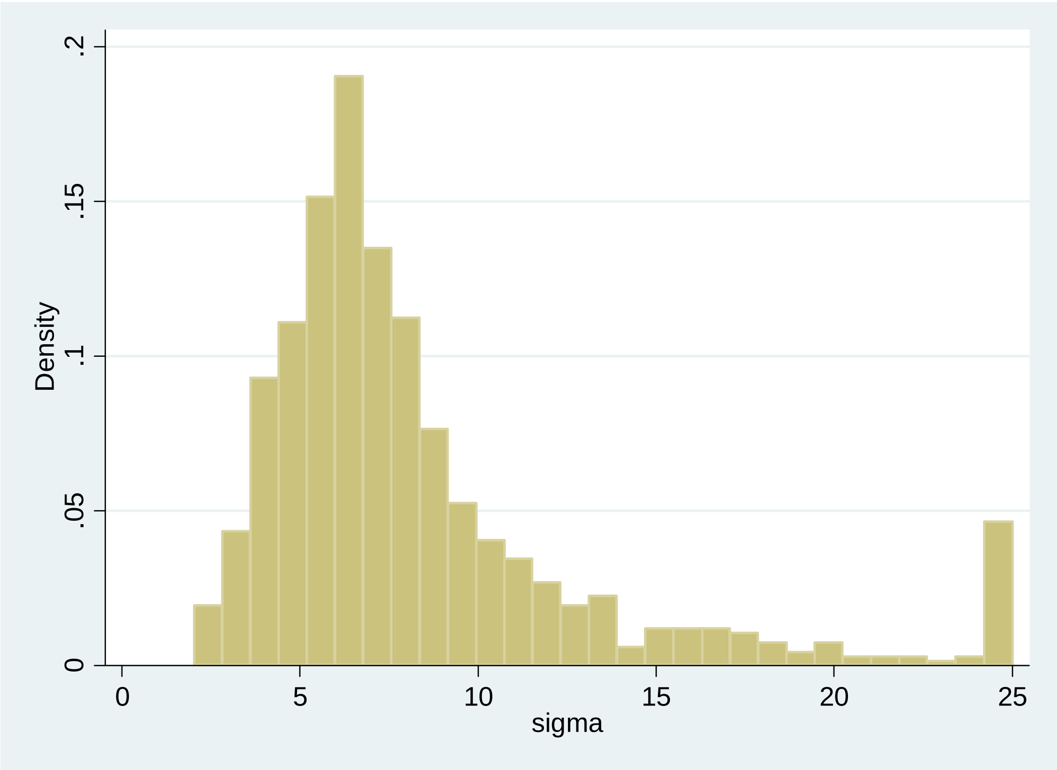

Figure 2: Estimates of γ

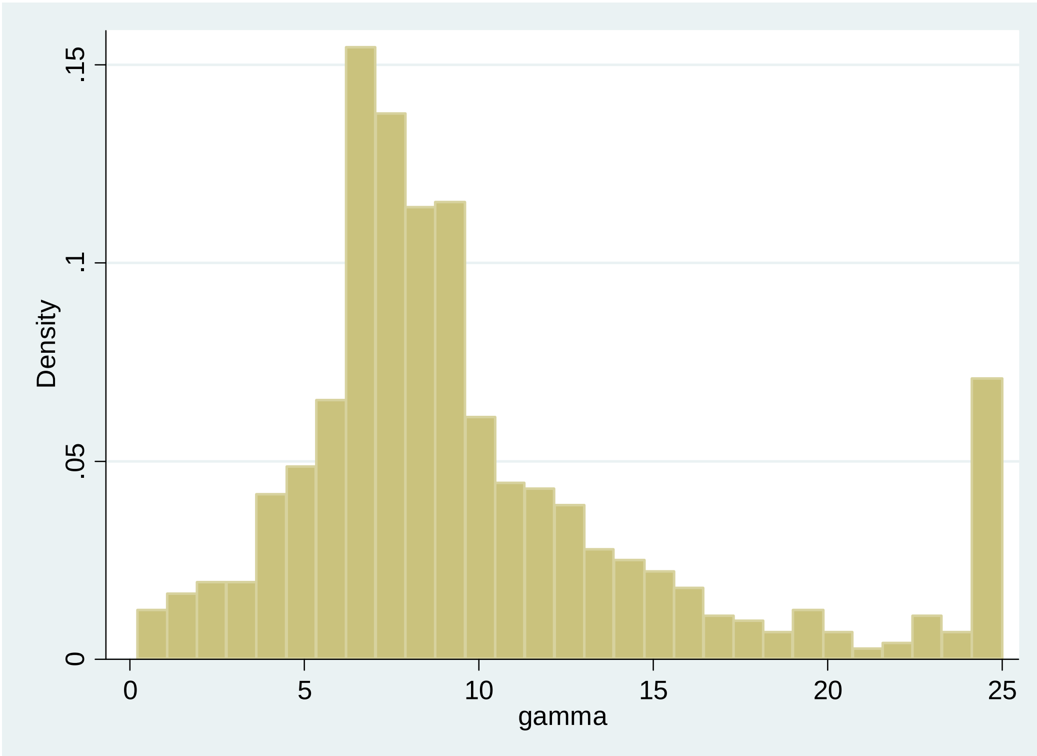

Figure 3: Estimates of θ

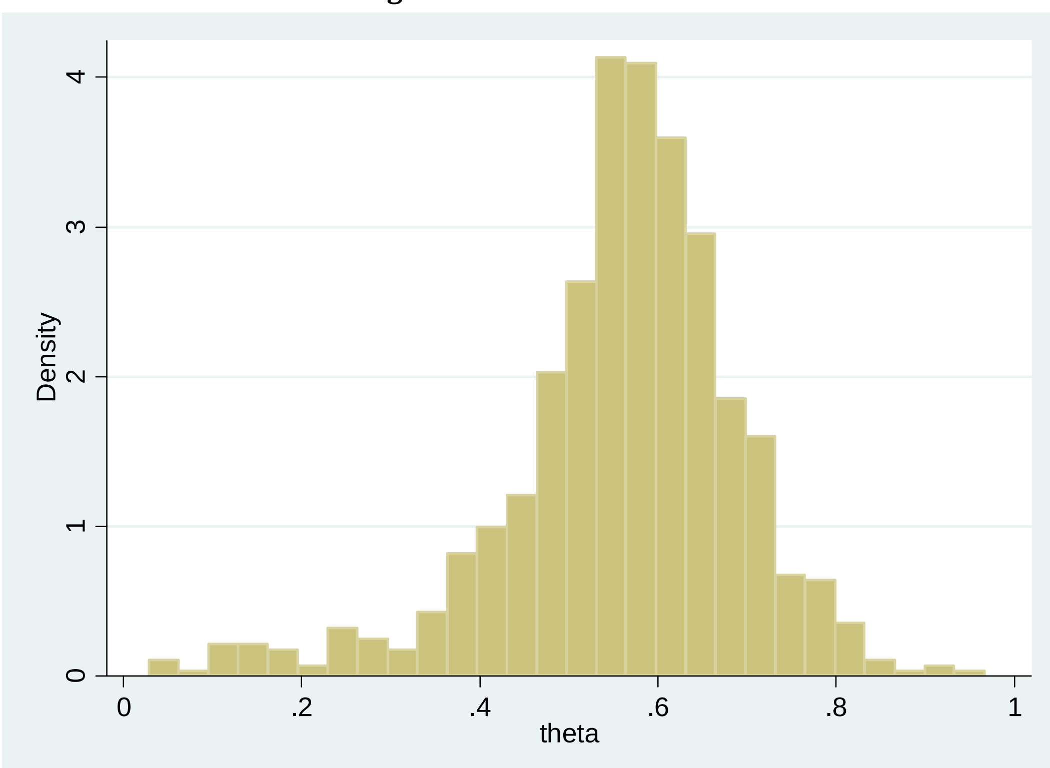

Table 2: Elasticity of Trade Quantity with Respect to Price
(estimated parameters σ, or γ+A+1 in the last row)

Estimation Sample and method:	Median Elasticity
USA Imports Only	2.7
+ All Bilateral Trade	4.7
+ Dropping Unreliable Observations	5.4
+ Estimating in Levels	8.0
+ Modeling Quality (using f.o.b. price)	14.7

bilateral trade rather than just imports into the United States. The second is that we are estimating a more complicated system, as seen from the estimating equation (24), where the elasticity of trade value with respect to the c.i.f. price is $\gamma+A$, whereas in Broda and Weinstein that elasticity is $(\sigma - 1)$. Both explanations have merit, as shown in Table 2.

In Table 2 we report estimates of the elasticity of trade *quantity* with respect to price, which corresponds to σ from Broda and Weinstein. In the first row we report the estimates using U.S. imports only, and like Broda and Weinstein estimate the model in first-differences *without* any quality adjustment, obtaining a similar median elasticity of 2.7. Extending our analysis to all bilateral trade further raises the median elasticity to 4.7. Since we have trade reports from both the exporting and the importing countries we can drop “unreliable” observations which are most subject to measurement or reporting error and are likely to attenuate elasticity estimates. Simply dropping the 5 percent of observations with the largest log-difference between reported unit values in the exporter's report and the importer's report raises our typical elasticity to 5.4. Estimation in levels rather than in differences raises our median estimate to 8.0, which suggests that elasticities are higher in the long-run than in the short-run, which may be partly captured by our levels estimation.[24] By these various steps, our expanded and filtered sample, in levels, has a much higher elasticity of substitution than found by Broda and Weinstein .

That is not the whole story, however, because while all these rows in Table 2 (except the last) use only the c.i.f. price in the specification of the demand equation, in fact, the f.o.b. price should also appear as a control for product quality as in (21).[25] Introducing the f.o.b. price, the total elasticity of trade value with respect to the c.i.f. price is the parameter $(\gamma + A)$ in (24), which has a median value of 13.7 in our estimates. This is converted to an elasticity of trade *quantity* with respect to the c.i.f. price of 14.7 and recorded in the last row of Table 2. That total elasticity combines both the response of trade on the intensive margin and the response on the extensive margin, and corresponds to median estimates for σ of 6.9 and γ of 8.4 in our sample. So besides the expanded sample, the other reason we find a greater responsiveness of trade is because we have used both margins in our specification.[26]

Our median estimate for the Pareto parameter γ of 8.4 is quite close to that reported by Eaton and Kortum (2002), who also considered bilateral trade between many countries.[27] Turning to θ, Crozet, Head and Meyer (2009) 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 lower than our median of 0.58 but well within the range of our parameter estimates shown in Figure 3.

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[24] Ideally, to estimate both short-run and long-run elasticities we would model the dynamics of demand and supply responses to price changes.

[25] Alternatively, the f.o.b. price can be interpreted as a control for the *ad valorem* trade costs in (22).

[26] Other authors, too, have found higher elasticities than Broda and Weinstein. Hummels (1999) directly estimates significant elasticities for many 2-digit SITC industries with an average value of 5.6 using OLS estimation and 9.3 using NLS (see his Tables 4 and 5). Erkel-Rousse and Mirza (2002) obtain pooled estimates for 27 ISIC industries of between 4 and 15. Romalis' (2007) NAFTA study estimates the average elasticity for HS 6-digit industries to be typically between 3 and 10, with estimates typically being between 6 and 10 when employing the most reliable (U.S.) tariff data

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

6. Indexes of Quality-Adjusted Prices and Quality

From (18) we can use the estimated parameters to construct the quality-adjusted prices, and (20) shows how to simplify this equation when comparing two exporters *i* and *j* selling to the same destination *k*. Alternatively, when comparing a given exporter *i* selling to two destinations *k* and *m*, we re-write (18) using (9) as,

$$\frac{\overline{\overline{P_{it}^k}}}{\overline{\overline{P_{it}^m}}} = \left(\frac{\overline{\overline{p_{it}^k}}}{\overline{\overline{p_{it}^m}}} \right)^{\rho} \left(\frac{Y_t^k \ / \ f_t^{k(1+B)} M_t^k}{Y_t^m \ / \ f_t^{m(1+B)} M_t^m} \right)^{\frac{\theta}{(1+\gamma)}}, \tag{30}$$

where $\rho \equiv (1-\theta)/[1+\theta(\sigma-1)]$. To implement this equation we need a solution for the market potential and the fixed costs in country *k*. Substituting (23) into (19), we can readily solve for the term $f_t^{kB} M_t^k$ using the estimated parameters of the model.[28] We further assume that the fixed costs f_t^k are equal across countries to measure (30). In future work, we could explore the suggestion of Arkolakis (2010) that fixed costs vary in proportion to the market size of the destination country.

To implement these formula for quality-adjusted prices we will rely on unit values. Re-writing (20) and (30) in terms of unit values, we actually measure:

$$\frac{UV_{it}^k}{UV_{jt}^k} \equiv \left(\frac{uv_{it}^k}{uv_{jt}^k} \right)^{\rho}, \qquad \rho \equiv \frac{(1-\theta)}{[1+\theta(\sigma-1)]}, \tag{31}$$

and,

$$\frac{UV_{it}^k}{UV_{it}^m} \equiv \left(\frac{uv_{it}^k}{uv_{it}^m} \right)^{\rho} \left(\frac{Y_t^k \ / \ f_t^{k(1+B)} M_t^k}{Y_t^m \ / \ f_t^{m(1+B)} M_t^m} \right)^{\frac{\theta}{(1+\gamma)}}. \tag{32}$$

[28] Note that $f_t^{kB} M_t^k$ equals the coefficient of a destination country fixed effect in (22), which is an equivalent way to solve for that term.

These unit values 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 EKS method, which is a many-country generalization of Fisher Ideal indexes, as we shall describe. [29] 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 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 in (31) 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 in question to $k=1,\ldots,C_{ij}$ destination markets. In a slight abuse of our earlier notation, let $uv_{it}^{*k} \equiv uv_{it}^{k} / tar_{it}^{k}$ denote the net-of-tariff unit value of exports to country *k*, but inclusive of transport costs.[30] Then the Laspeyres and Paasche price indexes of these export unit values are:

[29] Or the GEKS system, after Gini, Eltetö and Köves, and Szulc. We refer the reader to Balk (2008) for a modern treatment and complete 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 EKS generalization.

[30] We follow the convention in national accounts of measuring trade prices net of tariffs. We break from the convention of using f.o.b. prices for exports, however, by instead using c.i.f. unit values net of tariffs. This is done for consistency with the use of c.i.f. unit values in the quality-adjusted price formula (31) and (32).

$$P_{ijt}^{L} \equiv \frac{\sum_{k=1}^{C_{ij}} uv_{it}^{*k} q_{jt}^{k}}{\sum_{k=1}^{C_{ij}} uv_{jt}^{*k} q_{jt}^{k}}, \text{ and, } P_{ijt}^{A} \equiv \frac{\sum_{k=1}^{C_{ij}} uv_{it}^{*k} q_{it}^{k}}{\sum_{k=1}^{C_{ij}} uv_{jt}^{*k} q_{it}^{k}}. \tag{33}$$

In these expressions, q_{it}^{k} and q_{jt}^{k} are the quantity exported by countries *i* and *j* to country *k.*

Alternatively, we could instead we use the *quality-adjusted* unit values $UV_{it}^{*k} \equiv UV_{it}^{k} / tar_{it}^{k}$ in these formulas, in which case the quantities are instead Q_{it}^{k} with $uv_{it}^{*k} q_{it}^{k} = UV_{it}^{*k} Q_{it}^{k}$ 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. For example, letting $s_{jt}^{*k} = uv_{jt}^{*k} q_{jt}^{k} / \sum_{k} uv_{jt}^{*k} q_{jt}^{k}$ denote the export shares for country *j*, the Laspeyres index in (33) equals $P_{ijt}^{L} = \sum_{k} s_{jt}^{*k} (uv_{it}^{*k} / uv_{jt}^{*k})$. Likewise, the Paasche index is a weighted average of the unit-value ratios using the export shares s_{it}^{*k} of country *i*. In either case, we can alternatively use the ratio of quality-adjusted unit values, $(UV_{it}^{*k} / UV_{jt}^{*k})$, which is given by (31). 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, $P_{ijt}^{F} = (P_{ijt}^{L} P_{ijt}^{A})^{0.5}$. Then the EKS 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*:

$$P_{ikt}^{EKS} \equiv \prod_{j=1}^{C} \left(P_{ijt}^{F} P_{jkt}^{F} \right)^{1/C}, \tag{34}$$

with $P_{iit}^{F} \equiv 1$ for i =1,…,C. In most applications, the resulting EKS indexes are transitive.[31] That

[31] This is shown from (34) by noting that $P_{jkt}^{EKS} = 1 / P_{kjt}^{EKS}$, so that we readily compute $P_{ikt}^{EKS} P_{kmt}^{EKS} = P_{imt}^{EKS}$.

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 (34) is actually taken over only 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 EKS transformation of the Fisher Ideal indexes in (34) 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.[32]

This EKS aggregation is done for each 4-digit SITC product. The second stage is to aggregate over the SITC products, where we again use Fisher Ideal indexes – now computed by summing over products rather than over partner countries as in (33) – together with the EKS 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.

For convenience, we will refer to the EKS index of unit values as the "price index" and the EKS 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 category – to obtain the index of export quality.

Indexes for Import Prices and Quality

Our treatment of imports is analogous to our treatment of exports, so we only highlight

[32] 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.

the differences. In the first stage, the Laspeyres and Paasche indexes are computed by summing over source countries *i* that importers *k* and *m* both purchase from. So we compare the import prices of countries *k* and *m* 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 *m*. That is, the Laspeyres and Paasche indexes depend on $uv_{it}^{*k} / uv_{it}^{*m}$, or alternatively on the quality-adjusted unit value $UV_{it}^{*k} / UV_{it}^{*m}$, obtained by dividing (32) by tariffs $(tar_{it}^{k} / tar_{it}^{m})$. We then compute the Fisher Ideal indexes and perform the EKS transformation, resulting in an index of the import prices for country *k* relative to a base country *m* for each SITC product.[33] In the second stage, we aggregate over products to obtain indexes of imports 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 **[THESE ARE PRELIMINARY]**

Figures 4 to 6 summarize our results for our 2005 benchmark year. We aggregate raw export prices for each 4-digit SITC product into an aggregate price index for exports, with the U.S. export price in 2005 normalized to 1. We then similarly aggregate our quality estimates, and plot these indexes for 2005 for all 166 countries then in the PWT in Figure 4. 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).

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

Figure 4: Exports - Raw Prices, Quality, and Quality Adjusted Prices in 2005

Figure 5: Relative Import Unit Values, Quality, and Quality Adjusted Prices in 2005

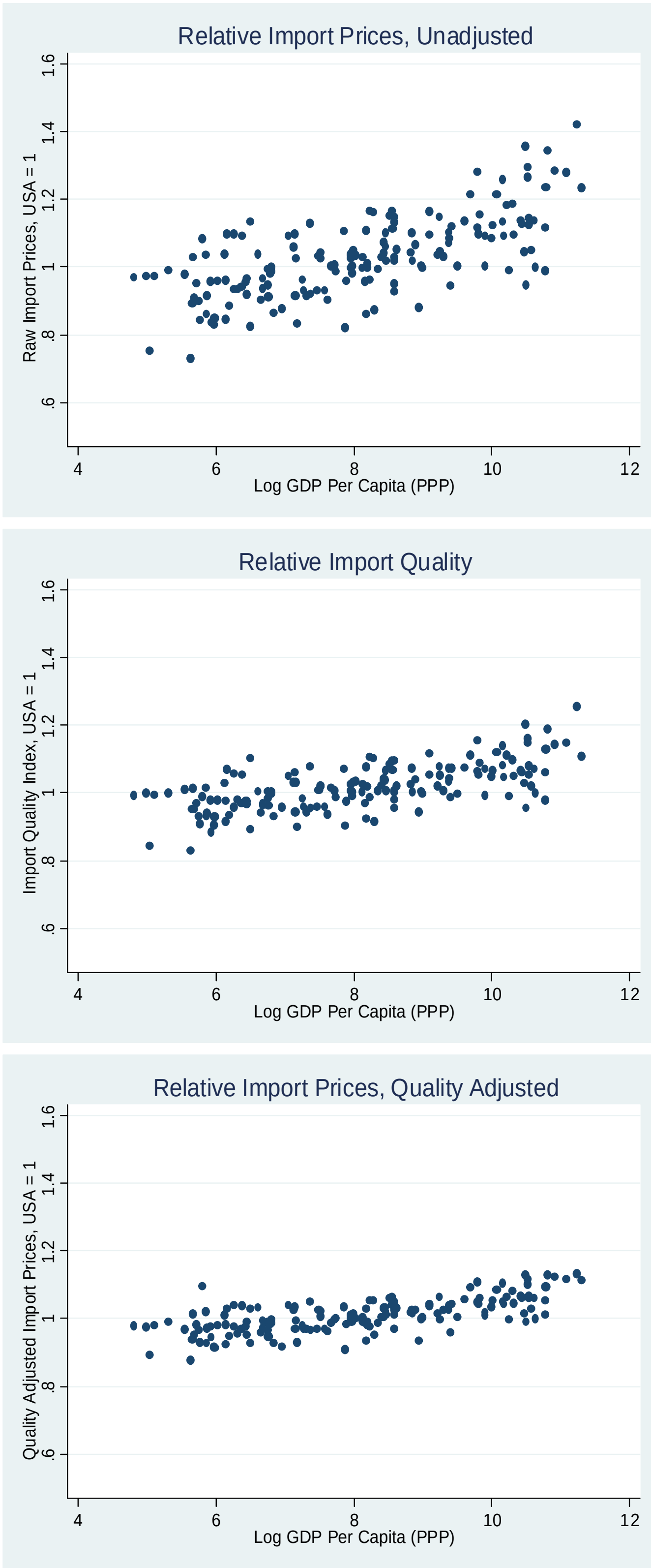
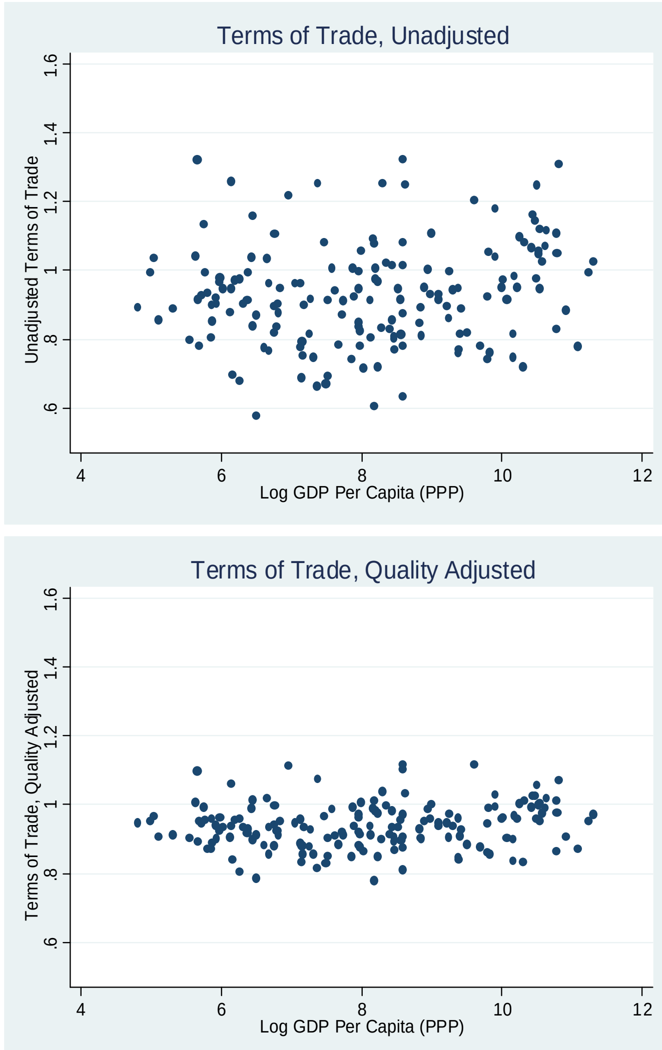


Figure 6: Terms of Trade 2005 - Unadjusted and Quality Adjusted



We illustrate a similar exercise for import prices in Figure 5. Developed countries import more expensive items (top panel) that are of higher quality (second panel). Quality-adjusted import prices increase moderately with the importing country's GDP per capita.

Figure 6 shows terms of trade estimates for our 2005 benchmark year. Terms of trade estimates constructed using raw export and import prices fluctuate substantially across countries, and lie between 0.57 and 1.32. Terms of trade estimates constructed from quality-adjusted prices move in a much narrower band, between 0.77 and 1.12. Despite the narrowness of this band, these quality-adjusted terms of trade measures are sufficiently different from 1 to produce meaningful differences between output-based and expenditure-based real GDP estimates for many countries in the PWT, even in our benchmark year.

We tabulate estimates for aggregate export quality for 1987, 1997 and 2007 in Table 3. This table includes the 50 largest countries ranked by the value of exports from 1984 to 2007, to which we add two other populous developing countries that also export many manufactures: Vietnam and Pakistan. Swiss exports have the highest quality, on average 41% higher than the U.S. in 2007, followed by Israel with quality 12 percent higher than the US. Japan and wealthy European countries also tend to have higher export quality than the US, though only moderately. Of note are the recent relatively high quality estimates for Mexico and for several Eastern European countries that have joined the EU, with aggregate export quality often within 10 percent of U.S. levels. Increased integration into the wealthy North American and Western European markets may be the root cause of this improvement. Asian countries that rapidly industrialized in the 1970's - such as Hong Kong, Singapore, South Korea and Taiwan - still lag U.S. export quality, with aggregate export quality usually between 10 an 20 percent lower than the US. Poorer large Asian countries have notably lower quality, with Indian and Chinese export quality respectively 23 percent and 34 percent lower than U.S. levels. Pakistan, Vietnam and Indonesia do little better, with quality lagging U.S. levels in 2007 by 21, 21 and 27 percent respectively. Also of interest is that China's relative export quality does not appear to have moved much despite substantial economic progress. This does not imply that its *absolute* export quality has not improved, but it's substantial exports of relatively low-quality products may have driven other countries to focus on even higher quality goods; see Amiti and Khandelwal (2009) for a related discussion.

Table 3: Export Quality in 1987, 1997 and 2007.
Quality Rankings, All Goods

	Rank				Normalized Quality, USA = 1			
country	1987	1997	2007	Change	1987	1997	2007	Change
Switzerland	1	1	1	0	1.32	1.44	1.41	0.09
Israel	15	14	2	13	0.97	1.05	1.12	0.15
Ireland	5	3	3	2	1.05	1.22	1.10	0.05
Austria	2	7	4	-2	1.09	1.14	1.06	-0.03
United Kingdom	9	6	5	4	1.01	1.17	1.05	0.04
Finland	6	8	6	0	1.04	1.13	1.05	0.01
Japan	4	5	7	-3	1.06	1.17	1.05	-0.01
Denmark	11	4	8	3	1.00	1.18	1.04	0.04
France	13	11	9	4	0.99	1.09	1.04	0.05
Sweden	7	2	10	-3	1.03	1.22	1.03	0.00
Germany	10	9	11	-1	1.01	1.10	1.01	0.01
USA	12	23	12	0	1.00	1.00	1.00	0.00
New Zealand	3	12	13	-10	1.08	1.08	1.00	-0.08
Belgium	16	19	14	2	0.96	1.01	1.00	0.04
Mexico	37	37	15	22	0.81	0.88	0.98	0.17
Canada	18	21	16	2	0.93	1.01	0.98	0.05
Italy	17	13	17	0	0.95	1.06	0.97	0.02
Australia	23	15	18	5	0.89	1.04	0.96	0.07
Hungary	44	28	19	25	0.77	0.94	0.95	0.18
Portugal	20	17	20	0	0.91	1.03	0.95	0.04
Netherlands	22	20	21	1	0.90	1.01	0.95	0.05
Czech Rep.		35	22	13		0.89	0.94	0.05
Spain	27	16	23	4	0.87	1.03	0.94	0.07
Norway	14	10	24	-10	0.98	1.09	0.94	-0.04
Chile	31	18	25	6	0.84	1.03	0.94	0.10
Singapore	26	26	26	0	0.87	0.95	0.91	0.04
Slovakia		39	27	12		0.87	0.88	0.02
South Africa	32	36	28	4	0.84	0.88	0.88	0.04
Philippines	39	33	29	10	0.80	0.90	0.87	0.07
Colombia	21	24	30	-9	0.91	0.98	0.87	-0.04
Poland	48	32	31	17	0.69	0.91	0.87	0.18
Korea, Rep.	28	22	32	-4	0.86	1.01	0.87	0.01
Algeria	29	27	33	-4	0.85	0.94	0.87	0.01
Thailand	47	40	34	13	0.72	0.86	0.86	0.13
Nigeria	35	34	35	0	0.84	0.89	0.85	0.02
Argentina	36	29	36	0	0.82	0.93	0.84	0.01
Turkey	25	25	37	-12	0.88	0.95	0.83	-0.05
Taiwan	45	48	38	7	0.76	0.81	0.83	0.07
Hong Kong	42	49	39	3	0.79	0.78	0.83	0.04
UAE	24	31	40	-16	0.88	0.92	0.83	-0.05
Brazil	41	30	41	0	0.80	0.92	0.82	0.02
Malaysia	38	50	42	-4	0.80	0.77	0.81	0.01
Russia		46	43	3		0.82	0.80	-0.01
Saudi Arabia	19	38	44	-25	0.91	0.87	0.80	-0.11
Pakistan	40	43	45	-5	0.80	0.85	0.79	-0.01
Viet Nam	30	41	46	-16	0.85	0.86	0.79	-0.06
India	33	47	47	-14	0.84	0.81	0.77	-0.07
Ukraine		51	48	3		0.74	0.77	0.03
Iran	8	45	49	-41	1.02	0.83	0.76	-0.26
Venezuela	43	42	50	-7	0.79	0.86	0.74	-0.05
Indonesia	34	44	51	-17	0.84	0.85	0.73	-0.11
China	46	52	52	-6	0.73	0.68	0.66	-0.07
Mean:					0.90	0.97	0.92	0.01
Standard Deviation:					0.12	0.14	0.13	0.01

Tables 4 through 9 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, usually higher than the US. Poor countries tend to have notably lower quality. Even the dispersion of the quality indexes is similar, save for capital goods and non-food consumer goods, where there is noticeably higher dispersion. Notable exceptions are in Table 6 for BEC 3: Fuels and Lubricants - where there is a less clear relationship between export quality and the exporter's level of development.

Some curious results in BEC 3: Fuels and Lubricants lead us to peer into the detail of the 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 using the Fisher Ideal Indexes. By 2007 Oman ceased to export SITC 3341 and its relative quality estimate in SITC 3345 had declined. Perhaps the main change in BEC 3 however is the relative rise in U.S. export quality, which is masked somewhat by our choice to normalize on U.S. quality. Peering into the underlying data shows that this is driven by a rise in the relative U.S. export quality of SITC 3330K ("Petroleum Oils and Crude Oils"), which is generated by many countries reporting small volumes of imports from the USA with high unit values. This high estimated quality of U.S. oil exports depresses the relative quality estimates in BEC 3 for any country whose dominant energy export is crude oil. The recent fairly high quality for Mexico shows up as an improvement in every BEC 1-digit

Table 4: Export Quality in 1987, 1997 and 2007.
Quality Rankings, BEC 1: Food and Beverages

	Rank				Normalized Quality, USA = 1			
country	1987	1997	2007	Change	1987	1997	2007	Change
Ireland	2	1	1	1	1.14	1.23	1.17	0.03
Australia	10	5	2	8	1.01	1.12	1.16	0.16
France	5	6	3	2	1.08	1.11	1.15	0.07
United Kingdom	4	3	4	0	1.10	1.17	1.14	0.05
New Zealand	1	4	5	-4	1.14	1.16	1.07	-0.07
Denmark	3	2	6	-3	1.10	1.19	1.05	-0.06
Germany	7	13	7	0	1.03	0.98	1.04	0.02
Netherlands	12	10	8	4	0.99	1.04	1.04	0.05
Italy	9	8	9	0	1.02	1.05	1.03	0.01
Belgium	6	7	10	-4	1.06	1.10	1.03	-0.04
Canada	8	9	11	-3	1.02	1.04	1.00	-0.02
USA	11	12	12	-1	1.00	1.00	1.00	0.00
Spain	13	11	13	0	0.96	1.02	1.00	0.04
Mexico	16	19	14	2	0.88	0.87	0.98	0.10
Malaysia	18	14	15	3	0.82	0.97	0.91	0.09
Brazil	15	16	16	-1	0.91	0.97	0.89	-0.02
Thailand	20	15	17	3	0.73	0.97	0.88	0.15
Indonesia	14	17	18	-4	0.94	0.93	0.86	-0.08
Argentina	19	20	19	0	0.79	0.84	0.83	0.04
China	17	18	20	-3	0.83	0.93	0.81	-0.03
Mean:					0.98	1.03	1.00	
Standard Deviation:					0.12	0.11	0.11	

Table 5: Export Quality in 1987, 1997 and 2007.
Quality Rankings, BEC 2: Industrial Supplies

	Rank				Normalized Quality, USA = 1			
country	1987	1997	2007	Change	1987	1997	2007	Change
Switzerland	1	1	1	0	1.40	1.39	1.24	-0.16
Japan	2	2	2	0	1.22	1.24	1.20	-0.02
United Kingdom	5	3	3	2	1.08	1.13	1.08	0.00
Hong Kong	13	11	4	9	0.93	0.99	1.06	0.13
Austria	3	5	5	-2	1.09	1.06	1.05	-0.04
France	7	8	6	1	1.01	1.00	1.02	0.01
Sweden	4	4	7	-3	1.08	1.09	1.01	-0.07
USA	9	9	8	1	1.00	1.00	1.00	0.00
Germany	8	10	9	-1	1.01	0.99	0.99	-0.02
Italy	6	6	10	-4	1.02	1.03	0.97	-0.05
Canada	11	12	11	0	0.94	0.96	0.97	0.02
Australia	17	14	12	5	0.89	0.95	0.95	0.05
Korea, Rep.	10	7	13	-3	0.95	1.01	0.95	0.00
Netherlands	16	16	14	2	0.92	0.90	0.94	0.02
Belgium	14	17	15	-1	0.93	0.90	0.93	0.00
Spain	18	15	16	2	0.86	0.91	0.93	0.07
Taiwan	12	13	17	-5	0.93	0.95	0.90	-0.04
Brazil	19	18	18	1	0.83	0.87	0.83	0.01
Russia		19	19	0		0.77	0.83	0.06
China	20	20	20	0	0.77	0.75	0.69	-0.08
Mean:					0.99	0.99	0.98	
Standard Deviation:					0.14	0.15	0.12	

Table 6: Export Quality in 1987, 1997 and 2007.
Quality Rankings, BEC 3: Fuels and Lubricants

	Rank				Normalized Quality, USA = 1			
country	1987	1997	2007	Change	1987	1997	2007	Change
Iraq	18	16	1	17	0.89	0.85	1.07	0.18
USA	6	4	2	4	1.00	1.00	1.00	0.00
Australia	11	8	3	8	0.97	0.92	0.96	-0.01
Mexico	19	20	4	15	0.88	0.78	0.95	0.07
United Kingdom	3	5	5	-2	1.04	0.98	0.94	-0.10
Netherlands	5	10	6	-1	1.01	0.92	0.89	-0.11
Oman	1	2	7	-6	1.25	1.07	0.89	-0.36
Saudi Arabia	9	12	8	1	0.97	0.88	0.86	-0.11
United Arab Emirate	13	11	9	4	0.95	0.89	0.85	-0.10
Algeria	14	19	10	4	0.93	0.80	0.84	-0.09
Nigeria	8	3	11	-3	0.99	1.01	0.84	-0.15
Venezuela	16	7	12	4	0.91	0.95	0.81	-0.09
Malaysia	2	13	13	-11	1.05	0.88	0.80	-0.25
Canada	15	9	14	1	0.91	0.92	0.79	-0.11
Kuwait	7	18	15	-8	1.00	0.81	0.76	-0.23
Iran	4	15	16	-12	1.02	0.86	0.74	-0.28
Qatar	17	1	17	0	0.90	1.13	0.72	-0.17
Norway	12	6	18	-6	0.96	0.96	0.72	-0.24
Russia		17	19	-2		0.82	0.71	-0.11
Indonesia	10	14	20	-10	0.97	0.86	0.67	-0.30
Mean:					0.98	0.91	0.84	
Standard Deviation:					0.08	0.09	0.11	

Table 7: Export Quality in 1987, 1997 and 2007.
Quality Rankings, BEC 4: Capital Goods and Parts

	Rank				Normalized Quality, USA = 1			
country	1987	1997	2007	Change	1987	1997	2007	Change
Switzerland	1	1	1	0	1.37	1.59	1.27	-0.11
Ireland	6	4	2	4	1.00	1.26	1.11	0.11
Canada	2	8	3	-1	1.07	1.10	1.10	0.03
Sweden	3	2	4	-1	1.05	1.34	1.05	0.00
United Kingdom	9	5	5	4	0.93	1.17	1.02	0.09
Germany	5	3	6	-1	1.01	1.33	1.01	0.00
USA	7	11	7	0	1.00	1.00	1.00	0.00
Belgium	10	6	8	2	0.89	1.13	1.00	0.10
Netherlands	8	10	9	-1	0.95	1.08	0.98	0.03
France	4	9	10	-6	1.01	1.09	0.95	-0.05
Mexico	16	15	11	5	0.74	0.77	0.94	0.20
Japan	11	7	12	-1	0.87	1.10	0.92	0.05
Italy	12	12	13	-1	0.83	0.96	0.88	0.05
Singapore	14	14	14	0	0.75	0.83	0.85	0.10
Malaysia	13	17	15	-2	0.78	0.70	0.85	0.06
Korea, Rep.	17	13	16	1	0.62	0.83	0.83	0.21
Thailand	15	16	17	-2	0.74	0.71	0.82	0.08
Taiwan	20	18	18	2	0.54	0.61	0.76	0.22
Hong Kong	18	19	19	-1	0.60	0.53	0.67	0.07
China	19	20	20	-1	0.55	0.44	0.56	0.01
Mean:					0.87	0.98	0.93	
Standard Deviation:					0.21	0.30	0.16	

Table 8: Export Quality in 1987, 1997 and 2007.
Quality Rankings, BEC 5: Transport Equipment and Parts

	Rank				Normalized Quality, USA = 1			
country	1987	1997	2007	Change	1987	1997	2007	Change
Austria	4	3	1	3	0.94	1.09	1.09	0.15
Canada	5	9	2	3	0.91	0.95	1.05	0.14
Belgium	7	7	3	4	0.86	1.00	1.04	0.18
United Kingdom	6	2	4	2	0.88	1.11	1.04	0.16
Hungary	17	14	5	12	0.63	0.93	1.03	0.40
Sweden	1	1	6	-5	1.01	1.15	1.02	0.01
Germany	3	4	7	-4	0.96	1.08	1.01	0.05
France	8	5	8	0	0.86	1.03	1.01	0.15
USA	2	6	9	-7	1.00	1.00	1.00	0.00
Mexico	14	13	10	4	0.73	0.93	0.98	0.25
Italy	9	11	11	-2	0.85	0.94	0.96	0.11
Czech Rep.		16	12	4		0.83	0.96	0.13
Netherlands	10	8	13	-3	0.85	0.97	0.96	0.11
Japan	11	12	14	-3	0.81	0.94	0.96	0.15
Spain	12	10	15	-3	0.74	0.95	0.95	0.22
Brazil	15	15	16	-1	0.71	0.91	0.91	0.21
Poland	19	18	17	2	0.56	0.78	0.88	0.32
Taiwan	16	19	18	-2	0.67	0.71	0.81	0.14
Korea, Rep.	13	17	19	-6	0.73	0.83	0.78	0.05
China	18	20	20	-2	0.58	0.60	0.69	0.11
Mean:					0.80	0.94	0.96	
Standard Deviation:					0.14	0.14	0.10	

Table 9: Export Quality in 1987, 1997 and 2007.
Quality Rankings, BEC 6: Consumer Goods

	Rank				Normalized Quality, USA = 1			
country	1987	1997	2007	Change	1987	1997	2007	Change
Switzerland	1	1	1	0	1.48	1.73	1.56	0.08
Japan	6	2	2	4	1.11	1.37	1.23	0.12
Ireland	3	3	3	0	1.16	1.29	1.08	-0.09
France	2	5	4	-2	1.21	1.28	1.07	-0.13
United Kingdom	7	7	5	2	1.06	1.19	1.07	0.00
Italy	5	6	6	-1	1.12	1.24	1.07	-0.05
Canada	10	12	7	3	0.98	0.96	1.04	0.06
Belgium	8	9	8	0	1.01	1.15	1.03	0.02
USA	9	11	9	0	1.00	1.00	1.00	0.00
Germany	4	4	10	-6	1.12	1.28	1.00	-0.13
Singapore	14	14	11	3	0.85	0.89	0.97	0.13
Netherlands	11	8	12	-1	0.97	1.15	0.97	0.00
Spain	12	10	13	-1	0.91	1.05	0.91	0.00
Mexico	19	16	14	5	0.72	0.85	0.90	0.18
Hong Kong	17	19	15	2	0.75	0.72	0.90	0.15
Korea, Rep.	15	15	16	-1	0.80	0.88	0.84	0.04
Turkey	13	13	17	-4	0.87	0.89	0.78	-0.08
Taiwan	18	17	18	0	0.74	0.77	0.72	-0.01
India	16	18	19	-3	0.79	0.73	0.72	-0.07
China	20	20	20	0	0.62	0.63	0.62	0.00
Mean:					0.96	1.05	0.97	
Standard Deviation:					0.21	0.27	0.20	

category, while the improvement in Eastern Europe is very apparent in their transport equipment exports. China's relative quality has only risen in transport equipment and, in more recent years, capital goods.

Our estimates call out for a comparison with the quality estimates of Hallak and Schott (2011). We do this in Figures 7 and 8 using data from Table IV of Hallak and Scott, and logs of our Table 3 results to make them comparable. Figure 7 compares normalized quality levels in 1997 for the countries common to the two tables.[34] The correlation is extremely high, at 0.68. What is most striking about the comparison is the different variances of the two sets of estimates. The standard deviation of the Hallak-Schott quality estimates is 0.45, three times the standard deviation of our estimates. We are unsure if this is due to the way Hallak and Schott aggregate their product-level results, or whether it is a feature of the first and/or second stages of their quality estimation procedure. Some of the lower variance is almost certainly because our estimates are derived from global trade, and not just U.S. imports. Figure 8 compares long-differences of the quality estimates. There is no relationship between the two. The different "volatility" of the two sets of estimates is again striking. Our results suggest that aggregate quality moves quite slowly for most countries, but large changes in quality are quite common in Hallak-Schott. Across countries, the standard deviation of these quality changes is 0.49 in Hallak-Schott, six times higher than the 0.08 for our estimates.

8. Conclusion

Our goal to adjust observed trade unit values for quality to estimate trade prices. Trade price estimates will be used to construct an output-based measure of real GDP in the Penn World

[34] 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 7: Levels Comparison With Hallak and Schott (2011)

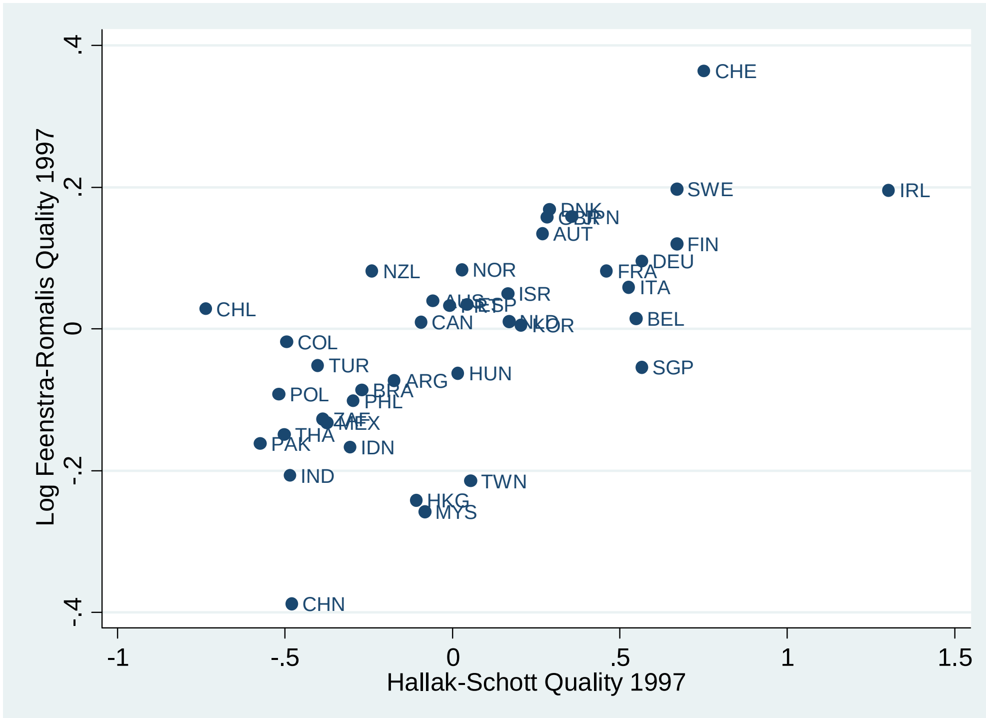


Figure 8: Differences Comparison With Hallak and Schott (2011)

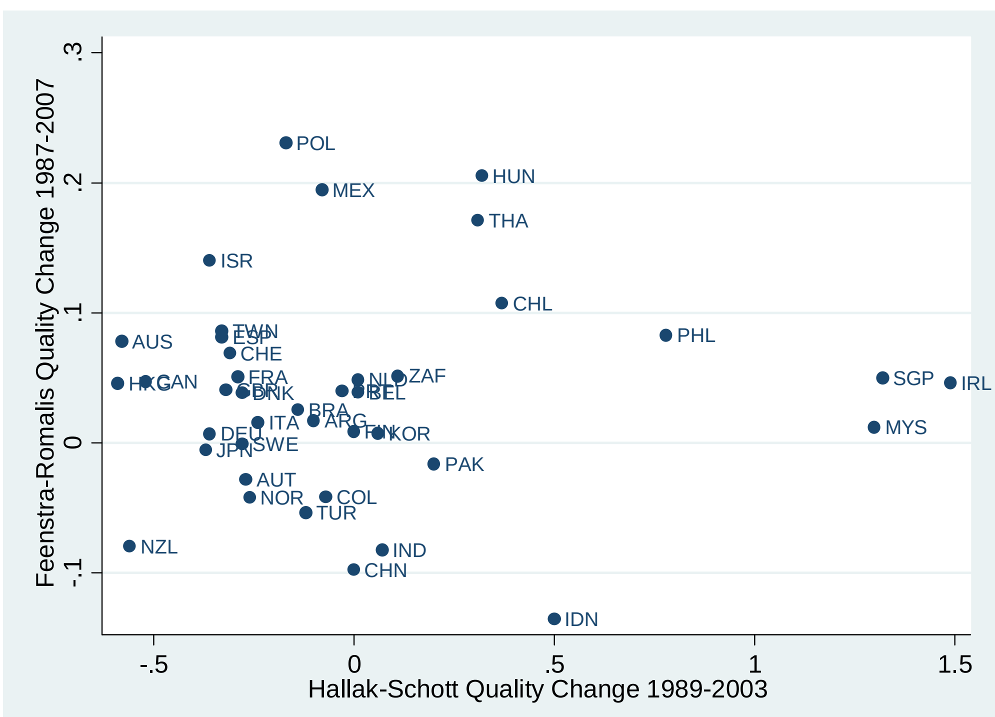


Table. We achieve this result by explicitly modeling quality-choice by exporting firms. Our key parameter estimate of the elasticity of quality with respect to the quantity of inputs almost always lies between 0 and 1, as required by our model. Our estimates of the elasticity of substitution between different varieties of the same SITC 4-digit products are substantially higher than in Broda-Weinstein, and the differences are large enough to warrant closer scrutiny. We reconcile our estimates with BW and account for why we get higher estimates. Finally, we use our estimates to construct quality and trade price estimates that we aggregate to the BEC 1-digit level and Total Export/Import level.

Appendix A:

We solve for the CES price index, which is:

$$P^k = \left[\sum_i M_i \overline{P_i^k}^{(1-\sigma)} \left(\frac{\hat{\varphi}_i^k}{\varphi_i} \right)^{-\gamma} \right]^{1/(1-\sigma)} . \tag{A1}$$

First, the quality-adjusted prices appearing in (A1) are replaced by using their solution in (17). Second, that solution from (17) is substituted back into (16) to obtain exports,[35]

$$\left(\frac{X_i^k / \tau_i^k f^k}{M_i (\varphi_i / w_i)^{\gamma}} \right) = \left(\frac{\overline{p_i^k}}{\overline{p_i^{*k}}^{\theta}} \right)^{-\frac{(\sigma-1)(1+\gamma)}{[1+\theta(\sigma-1)]}} \left(\frac{Y^k P^{k(\sigma-1)}}{\tau_i^k f^k} \right)^{\frac{(1+\gamma)}{[1+\theta(\sigma-1)]}} \kappa_5 . \tag{A2}$$

Third, we solve for the export probabilities $(\hat{\varphi}_i^k / \varphi_i)^{-\gamma}$ appearing in (A1) using (13),

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

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 (A1) to solve for the prices index in terms of those other variables. That solution is:

$$Y^k P^{k(\sigma-1)} = \left(\frac{Y^k}{M^k \kappa_5} \right)^{\frac{[1+\theta(\sigma-1)]}{(1+\gamma)}} ,$$

where M^k is defined by (19). Using this solution for the price index in (A2), we immediately obtain the gravity equation shown in (22).

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[35] $\kappa_5 = \{[\gamma - \theta(\sigma - 1)] / \sigma\gamma]\}^{\gamma/[1+\theta(\sigma-1)]} \kappa_1^{\gamma(\sigma-1)/[1+\theta(\sigma-1)]}$.

Appendix B:

(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[36]; (iii) U.S. tariff schedules from the U.S. International Trade Commission from 1989 onwards[37]; (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 US, 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

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

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 member (Norway). We aggregate MFN and each non-MFN tariff program[38] to the 4-digit SITC Revision 2 level by taking the simple average of tariff lines within each SITC code.

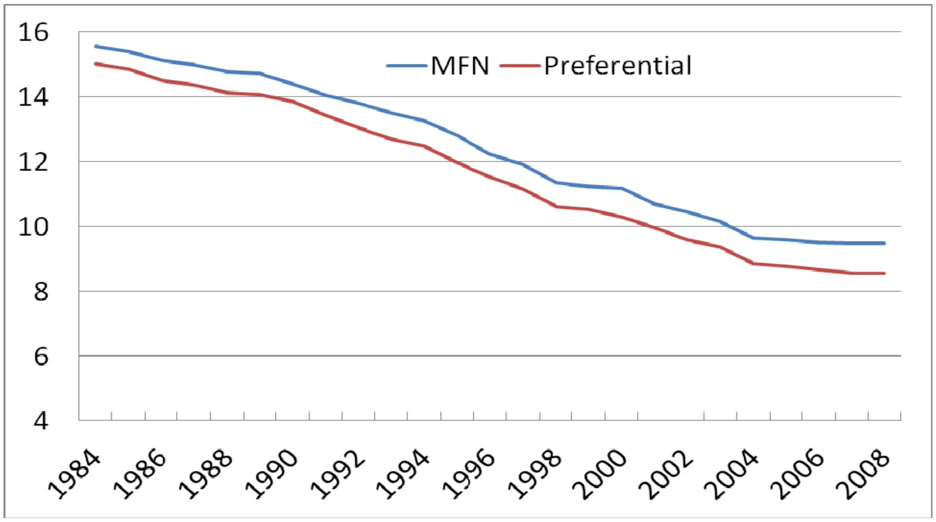
Tariff schedules are often not available in each year, especially for smaller countries. Updated schedules are more likely to be available after significant tariff changes. Rather than replacing "missing" MFN tariffs by linearly interpolating observations, missing observations are set equal to the nearest 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.

[38] 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.

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

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.

(iv) Quality-Adjusted Unit Values

The quality-adjusted prices shown by (31) and (32) depend on the c.i.f. prices. In our estimation we will be using both c.i.f. and f.o.b. unit values, but two-thirds of the bilateral, 4-digit SITC trade flows in our 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 (9) the ratio of c.i.f. to f.o.b. prices is $\overline{p_{it}^{k}} \big/ \overline{p_{it}^{*k}} = \kappa_4 \tau_{it}^{k}$, where $\kappa_4 \equiv \left[\left(\frac{1}{1-\theta}\right)\left(\frac{\sigma}{\sigma-1}\right)\right] \Big/ \left[\left(\frac{1}{1-\theta}\right)\left(\frac{\sigma}{\sigma-1}\right)-1\right]$. We use the estimates from the preliminary regression:

$$\ln(p_{it}^{k} / p_{it}^{*k}) = \beta_0 + \ln tar_{it}^{k} + \beta_1 \ln dist_i^{k} + \delta_{1it}^{k},$$

to form an estimate of the predicted *ad valorem* trade costs, $\widehat{\kappa_4\tau}_{it}^{k}$. Accordingly, 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 $uv_{it}^{*k} \times \widehat{\kappa_4\tau}_{it}^{k}$. Then the ratio of quality-adjusted unit-values is measured by either (31) or (32) in the text.

Appendix C:

We re-write (24) and (29) slightly as:

$$\begin{aligned}(\gamma + A)\left(\ln \overline{p_{it}^{k}} - \ln \overline{p_{jt}^{k}}\right) - \gamma\left(\ln \overline{p_{it}^{*k}} - \ln \overline{p_{jt}^{*k}}\right) + \left(\ln X_{it}^{k} - \ln X_{jt}^{k}\right) \\ = \alpha_i - \alpha_j + \alpha(\ln L_{it} - \ln L_{jt}) + \varepsilon_{it}^{k} - \varepsilon_{jt}^{k},\end{aligned} \tag{24'}$$

$$\begin{aligned}(\gamma + A + \omega_2\rho)\left(\ln \overline{p_{it}^{k}} - \ln \overline{p_{jt}^{k}}\right) - \gamma\left(\ln \overline{p_{it}^{*k}} - \ln \overline{p_{jt}^{*k}}\right) - \omega_2(\ln X_{it}^{k} - \ln X_{it}^{k}) \\ = \omega_0(\ln tar_{it}^{k} - \ln tar_{jt}^{k}) + \omega_1(dist_i^{k} - dist_j^{k}) + (\delta_{it}^{k} - \delta_{jt}^{k}).\end{aligned} \tag{29'}$$

Taking the product of (24') and (29') and dividing by $(\gamma + A)(\gamma + A + \omega_2\rho)$, we obtain:

$$\begin{aligned}\left(\ln \overline{p_{it}^{k}} - \ln \overline{p_{jt}^{k}}\right)^2 &= \frac{2\gamma(\gamma + A) + \gamma\omega_2\rho}{(\gamma + A)(\gamma + A + \omega_2\rho)}\left(\ln \overline{p_{it}^{k}} - \ln \overline{p_{jt}^{k}}\right)\left(\ln \overline{p_{it}^{*k}} - \ln \overline{p_{jt}^{*k}}\right) \\ &- \frac{\gamma^2}{(\gamma + A)(\gamma + A + \omega_2\rho)}\left(\ln \overline{p_{it}^{*k}} - \ln \overline{p_{jt}^{*k}}\right)^2 + \frac{\omega_2}{(\gamma + A)(\gamma + A + \omega_2\rho)}\left(\ln X_{it}^{k} - \ln X_{jt}^{k}\right)^2 \\ &+ \left(\frac{\omega_2}{(\gamma + A + \omega_2\rho)} - \frac{1}{(\gamma + A)}\right)\left(\ln X_{it}^{k} - \ln X_{jt}^{k}\right)\left(\ln \overline{p_{it}^{k}} - \ln \overline{p_{jt}^{k}}\right) \\ &+ \frac{\gamma(1 - \omega_2)}{(\gamma + A + \omega_2\rho)(\gamma + A)}\left(\ln X_{it}^{k} - \ln X_{jt}^{k}\right)\left(\ln \overline{p_{it}^{*k}} - \ln \overline{p_{jt}^{*k}}\right) \\ &+ \frac{(\alpha_i - \alpha_j)}{(\gamma + A + \omega_2\rho)(\gamma + A)}\left[\omega_0(\ln tar_{it}^{k} - \ln tar_{jt}^{k}) + \omega_1(dist_i^{k} - dist_j^{k})\right] \\ &+ \frac{\alpha(\ln L_{it} - \ln L_{jt})}{(\gamma + A + \omega_2\rho)(\gamma + A)}\left[\omega_0(\ln tar_{it}^{k} - \ln tar_{jt}^{k}) + \omega_1(dist_i^{k} - dist_j^{k})\right] + \mu_{it}^{k},\end{aligned} \tag{C1}$$

where the error term is:

$$\mu_{it}^k = \frac{(\delta_{it}^k - \delta_{jt}^k)}{(\gamma + A)(\gamma + A + \omega_2 \rho)} \left[\alpha_i - \alpha_j + \alpha(\ln L_{it} - \ln L_{jt}) + \varepsilon_{it}^k - \varepsilon_{jt}^k \right]$$
$$+ \frac{(\varepsilon_{it}^k - \varepsilon_{jt}^k)}{(\gamma + A)(\gamma + A + \omega_2 \rho)} \left[\omega_0 (\ln tar_{jt}^k - \ln tar_{it}^k) + \omega_1 (dist_i^k - dist_j^k) \right].$$

We treat the country fixed effects, distance and tariffs as exogenous, we can assume that they are uncorrelated with the demand and supply shocks. We further assume that the supply shocks are uncorrelated with the demand shocks, so that $E\mu_{it}^k = 0$ for each source country *i* and destination *k*. This is the moment condition we use to estimate (C1).

Notice from (20) and (25) we can solve for $\rho = A\zeta / [A(\zeta + \gamma) + \gamma(1+\gamma)]$. Substituting this into (C1) we obtain an equation that is nonlinear in the parameters γ, A and ω_2, with θ and σ then implied by (25). For estimation, we average the variables in (C1) 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 country fixed effects $(\alpha_i - \alpha_j)$ interacted with distance and tariffs as appear in (C1). 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 (C1) on those interaction terms, and then estimate (C1) using the residuals obtained from these preliminary regressions. In these first-stage regressions we also include the source and destination country fixed effects themselves. These are needed to control for measurement errors in the c.i.f. and f.o.b. unit values, which are used in place of prices in (C1). We model the measurement errors as in (26) and assume that they are independent of each other and of the import shares. Then the measurement errors will appear as variance terms in (C1) after averaging over time for a large enough number of periods. The country indicators then absorb these variances.

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