

International Prices and Endogenous Quality*

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Abstract

The unit values of internationally traded goods are heavily influenced by quality. We model this in an extended monopolistic competition framework where, in addition to choosing price, firms simultaneously choose quality. We allow countries to have non-homothetic demand for quality. The optimal choice of quality by firms reflects this non-homothetic demand as well as the costs of production, including specific transport costs as in the “Washington apples” effect. We estimate quality and quality-adjusted prices for 185 countries over 1984-2011. Our estimates are less sensitive to assumptions about the extensive margin than are “demand side” estimates. We find that quality-adjusted prices vary much less across countries than do unit values, and surprisingly, that the quality-adjusted terms of trade are negatively related to countries’ level of income.

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1. Introduction

The quality of internationally traded products has become an important area of study. Product quality is a key feature of how countries specialize in production (Schott, 2004), the direction of trade between countries (Hallak, 2006), and even how countries grow (Hummels and Klenow, 2005). Trade prices and countries' terms of trade have also long played a central role in international trade theory and international macroeconomics. Researchers studying these variables are often limited to statistics for individual nations, sometimes made available as short series in international databases such as the World Bank's World Development Indicators. This paper develops and implements a new methodology exploiting a pervasive supply-driven feature of trade data to decompose widely-available unit values of internationally traded goods into quality and quality-adjusted price components. Results for individual products for almost all countries from 1984-2011 are aggregated to industry-level indexes of import and export quality, import and export prices, and terms of trade.

We are not the first to attempt to disentangle quality from trade unit values, and other recent authors to do so include Schott (2004, 2008), Hallak (2006), Hallak and Schott (2011), Khandelwal (2010) and Martin and Méjean (2010).¹ These studies rely on the demand side to identify quality together with a simple supply-side to control for the extensive margin. In the words of Khandelwal (2010, p. 1451): “The 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, with higher net imports – conditional on price – implying higher quality.

¹ Another line of literature empirically distinguishes between productivity and quality versions of the Melitz (2003) model: see Baldwin and Ito (2011), Crozet, Head and Mayer (2012), Johnson (2012) and Mandel (2009).

To this demand-side intuition we will add a supply side, in two respects. First, our model of endogenous quality choice by firms, described in section 2, gives rise to a “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 quality and distance, or between exporter f.o.b. price and distance, is an immediate implication of the first-order condition of firms for optimal quality choice. It will allow us to use the exporter f.o.b. price to help identify quality.

We embed this quality decision into a Melitz (2003) model with heterogeneous firms, described in section 3. Included in the model is the zero-cutoff-profit condition that determines the marginal exporter. That condition is a second supply-side relation that will help us to identify quality, and it works in the opposite direction as the demand-side intuition. As foreign demand rises, less-efficient exporters enter and they produce lower quality. It follows that quality and bilateral trade are *negatively* related from this supply-side relation. Combined with the positive relationship between trade and quality from the demand side, we will obtain a much sharper solution for quality than previous literature. That solution depends on c.i.f. and f.o.b. prices (measured by unit values) and the parameters of our model: the elasticity of substitution, Pareto productivity parameter, and also a parameter governing non-homothetic demand, which we allow as in recent literature.²

In section 4, we estimate these parameters from a gravity-like equation implied by our model, using detailed bilateral trade data at the 4-digit SITC digit level (nearly 800 products per year) for 185 countries during 1984-2011. Our median estimate of the elasticity of substitution is higher than that in Broda and Weinstein (2006), which we attribute to several features: our

² Our specification of non-homothetic tastes is similar to that in Hallak (2006), but working with the expenditure rather than utility function.

expanded sample over many countries, the fact that quality is included, and by using a specification that is more robust to measurement error. Our median estimate of the Pareto parameter is quite close to the estimated Fréchet parameter in Eaton and Kortum (2002), who also consider trade between many countries.

Given the parameter estimates, product quality is readily constructed in section 5. On the export side we find that much of the variation in unit values is explained by quality, so that quality-adjusted prices vary much less than the raw unit values or than the quality-adjusted estimates of Hallak and Schott (2011), and Khandelwal (2010). We also find that our estimates are less sensitive to assumptions about the extensive margin than the “demand side” estimates of these authors (because we solve for the extensive margin in our model).

On the import side we find that quality-adjusted import prices tend to be lower for poor countries. It follows that countries' quality-adjusted terms of trade are negatively related to their level of income. This surprising result is due in part to the lower unit value of imports for poor countries, but also relies on the supply side of our model: countries with lower imports (because they are poor or just small) buy from more efficient foreign firms who can overcome the fixed costs of exporting, and these firms sell higher quality. Offsetting that effect is the reduced preference for quality in low-income countries. Balancing these opposing effects, import quality is only weakly related to country income. Since import unit values are more strongly related to income, it follows that the quality-adjusted import prices are lower for poor countries. This result lends support to the proposition of Fajgelbaum, Grossman and Helpman (2011a) that poorer countries are net importers of higher-quality goods (because they are not produced locally): we find that import quality is less related to income than is export quality, so that poorer countries do appear to be net importers of higher quality goods.

We provide indexes of quality and quality-adjusted prices for the 4-digit SITC and 1-digit Broad Economic Categories (distinguishing food and beverages, other consumer goods, capital, fuels, intermediate inputs and transport equipment), that should be useful to researchers interested in the time-series or cross-country properties of these indexes and that will be incorporated into the next generation of the Penn World Table (PWT; see Feenstra, Inklaar and Timmer, 2013). In addition to their use in PWT, the quality and price indexes produced by our study will find wide application in international trade and macroeconomics. For example, trade prices are important for the study of trade and wages (Lawrence and Slaughter, 1993). Capital goods prices are used in “development accounting” (Hsieh and Klenow, 2010). Intermediate goods prices are used to study the effects of trade on growth (Estevadeordal and Taylor, 2013). Terms of trade indexes are used to study the arguments for fixed versus flexible exchange rates (Broda, 2001) and the world income distribution (Acemoglu and Ventura, 2002). Finally, an extensive database of international tariffs collected for this paper will be useful for empirical international trade research.

2. Optimal Quality Choice

Consumer Problem

Consumers in country k have available a continuum i of differentiated varieties of a product in a sector. These products can come from different source countries. Denote the price and quality of good i in country k by p_i^k and z_i^k , respectively. Demand in country k arises from the expenditure function:

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

with, $\alpha^k = h(U^k) = 1 + \lambda \ln U^k$, for $U^k > 0$. (1b)

Quality z_i^k is raised to the power $\alpha^k > 0$, which we denote by $z_i^{\alpha^k} \equiv (z_i^k)^{\alpha^k}$ for brevity. Thus, quality acts as a shift parameter in the expenditure function. Hallak (2006) introduced a similar exponent on quality, but in the context of the *direct* utility function (as also used by Demir, 2012). In that case it is not possible to make the exponents α^k depend on utility or per-capita income; but by working with the expenditure function we are able to do just that. Because $\alpha^k = h(U^k)$ depends on utility, this expenditure function has non-homothetic demand for quality, as in Fajgelbaum, Grossman and Helpman (2011a,b).³

The assumptions of the CES functional form in (1a) and the parameterization of the exponents $h(U^k)$ in (1b) are both made for convenience. The key assumption is that *price is divided by quality in the expenditure function*, enabling us to reformulate consumer decisions in terms of *quality-adjusted* prices and quantities. Differentiating this expenditure function to compute demand q_i^k :

$$q_i^k = \frac{\partial E^k}{\partial p_i^k} = \frac{\partial E^k}{\partial P_i^k} \frac{1}{z_i^{\alpha^k}},$$

where we define the *quality-adjusted prices* $P_i^k \equiv p_i^k / z_i^{\alpha^k}$, which are the natural arguments of the expenditure function in (1a). Likewise defining *quality-adjusted demand* $Q_i^k \equiv z_i^{\alpha^k} q_i^k$, we can re-arrange terms above to obtain $Q_i^k = \partial E^k / \partial P_i^k$. It follows that working with the quality-adjusted magnitudes still gives quantity as the derivative of the expenditure function with respect to price.

The expenditure function in (1) is valid provided that it is increasing in utility and non-

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

decreasing in price.⁴ Using the assumed functional forms, we derive:

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

since $dP_i^k / dU^k = -P_i^k \ln z_i^k h'(U^k) = -\lambda P_i^k \ln z_i^k / U^k$, using $P_i^k \equiv p_i^k / z_i^{\alpha^k}$ and (1b). The final integral above is interpreted as the average of log quality across products. Thus, the expenditure function in (1) is increasing in utility provided that λ is sufficiently small, which is readily confirmed in our estimates.

Firms' Problem

The production side of the model is an extension of Melitz (2003) to allow for endogenous quality choice by firms. The detailed assumptions are as follows:

A1. *Firms may produce multiple products, one for each potential market.*

A2. *Firm j producing in country i simultaneously chooses the quality z_{ij}^k and free on board (f.o.b.) price p_{ij}^{*k} for each market k .*

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.⁵ This assumption allows for a convenient solution for quality and was used by Rodriguez (1979) and other early literature dealing the impact of import quotas on product quality.⁶ Much of the recent literature on product quality in trade also adopts assumption **A2** when quality is treated as

⁴ The idea of allowing the parameters of the expenditure function to depend on utility is borrowed from Deaton and Muellbauer (1980, pp. 154-158), who define an expenditure function as a utility-weighted combination of any two functions that are non-decreasing in price, which is valid provided that the resulting function is increasing in utility.

⁵ To justify our assumption that quality characteristics are changed just as often as prices, we can look to the example of the “voluntary” export restraint (VER) on Japanese auto exports to the United States in the early 1980s. As documented by Feenstra (1988), the characteristics of the Japanese exports were changed on the same annual basis as their prices.

⁶ For example, Krishna (1987) and Das and Donnenfeld (1987).

endogenous: see Mandel (2009), Khandelwal (2010), Antoniadis (2012), Demir (2012) and Johnson (2012, Appendix), for example.⁷

A3. To produce each unit of a good with quality z_{ij}^k , the firm with productivity φ_{ij} must use a composite input (“labor”) l_{ij}^k according to the Cobb-Douglas production function:

$$z_{ij}^k = (l_{ij}^k \varphi_{ij})^\theta, \quad (2)$$

where $0 < \theta < 1$ reflects diminishing returns to quality.

The Cobb-Douglas functional form in (2) is used for convenience, similar to Verhoogen (2008).

In later work, Kugler and Verhoogen (2012) have used a CES functional form. We discuss below and in Appendix A how assumption **A3** can be generalized while retaining the convenient log-linear results that we shall derive. This generalization would be challenging to implement for data reasons, however, so we rely on the Cobb-Douglas formulation in (2).

A4. Productivity is Pareto distributed with distribution function $G_i(\varphi) = 1 - (\varphi / \varphi_i)^{-\gamma}$, where the location parameter $\varphi_i \leq \varphi$ is 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, but for analytical convenience we assume that the dispersion parameter γ is identical across countries.⁸

A5. There are both specific trade costs T_i^k and ad valorem trade costs between countries i and k .

One plus the ad valorem trade costs are denoted by τ_i^k , which includes one plus the ad valorem tariff, denoted by tar_i^k . The ad valorem trade costs are applied to the value inclusive of the specific trade costs.⁹ The tariff-inclusive c.i.f. price therefore is $p_{ij}^k \equiv \tau_i^k (p_{ij}^{*k} + T_i^k)$, and the net-

⁷ Gervias (2010) has quality chosen for the lifetime of a product, so he does not use assumption A2.

⁸ In this respect we are making the same assumption as in Eaton and Kortum (2002), who allowed for different location parameters of the Fréchet distribution across countries, but with the same dispersion parameter.

of-tariff c.i.f. price is $p_{ij}^k / \text{tar}_i^k$.¹⁰

A6. Firms must pay fixed costs of $f_i^k(\varphi_{ij})$ to export, which depends on their productivity φ_{ij} .

We include a detailed discussion of the specification of fixed costs in section 3 of the paper.

We now solve for the optimal f.o.b. price p_{ij}^{*k} and quality z_{ij}^k that a firm simultaneously chooses for each destination market, conditional on exporting (in section 3 we will turn to the export decision). We denote the price of the composite input l_{ij}^k by the wage w_i . The marginal cost of producing a good of quality z_{ij}^k is then solved from (2) as,

$$c_{ij}(z_{ij}^k, w_i) = w_i l_{ij}^k = w_i (z_{ij}^k)^{1/\theta} / \varphi_{ij}. \quad (3)$$

From the iceberg costs, τ_i^k units of the good are exported in order for one unit to arrive, so total exports are $y_{ij}^k = \tau_i^k q_{ij}^k$. When evaluating profits from exporting to country k , we need to divide by one plus the *ad valorem* tariff tar_i^k , obtaining:

$$\begin{aligned} \max_{p_{ij}^{*k}, z_{ij}^k} [p_{ij}^{*k} - c_{ij}(z_{ij}^k, w_i)] \frac{\tau_i^k q_{ij}^k}{\text{tar}_i^k} &= \max_{p_{ij}^{*k}, z_{ij}^k} \left[\frac{p_{ij}^{*k}}{z_{ij}^{\alpha^k}} - \frac{c_{ij}(z_{ij}^k, w_i)}{z_{ij}^{\alpha^k}} \right] \frac{\tau_i^k Q_{ij}^k}{\text{tar}_i^k} \\ &= \max_{p_{ij}^k, z_{ij}^k} \left\{ p_{ij}^k - \tau_i^k \frac{[c_{ij}(z_{ij}^k, w_i) + T_i^k]}{z_{ij}^{\alpha^k}} \right\} \frac{Q_{ij}^k}{\text{tar}_i^k}. \end{aligned} \quad (4)$$

The first equality in (4) converts from observed to quality-adjusted consumption, while the

second line converts to quality-adjusted, tariff-inclusive c.i.f. prices $P_{ij}^k \equiv \tau_i^k (p_{ij}^{*k} + T_i^k) / z_{ij}^{\alpha^k}$,

⁹ Most countries apply tariffs to the transport-inclusive (c.i.f.) price of a product. The exceptions are Afghanistan, Australia, Botswana, Canada, Democratic Republic of the Congo, Lesotho, Namibia, New Zealand, Puerto Rico, South Africa, Swaziland, and the United States. See the Customs Info Database at <http://export.customsinfo.com/> and http://export.gov/logistics/eg_main_018142.asp. If we instead assumed that *ad valorem* trade costs only applied to the f.o.b. price, then we would replace T_i^k with T_i^k / τ_i^k in our formulas below.

¹⁰ In our estimation we further model the costs as depending on distance and the quantity shipped, with the full specification in Appendix E.

while changing the choice variables from p_{ij}^{*k}, z_{ij}^k to P_{ij}^k, z_{ij}^k . This change in variables relies on our assumption **A2** that prices and characteristics are chosen simultaneously, but (4) does not rely on the functional forms in (1).

It is immediate that to maximize profits in (4), firms must choose z_{ij}^k to minimize $[c_{ij}(z_{ij}^k, w_i) + T_i^k] / z_{ij}^{\alpha^k}$. In the case where $\alpha^k = 1$, this problem is interpreted as *minimizing the average variable cost per unit of quality, inclusive of specific trade costs*, which is obtained where marginal cost equals average cost as found by Rodriguez (1979). More generally, with $\alpha^k > 0$ the solution to this problem is:

$$\frac{\partial c_{ij}(z_{ij}^k, w_i)}{\partial z_{ij}^k} = \alpha^k \frac{[c_{ij}(z_{ij}^k, w_i) + T_i^k]}{z_{ij}^k}, \quad (5)$$

so there is a wedge of α^k between the marginal and average costs of producing quality. The second-order condition for this minimization problem is satisfied if and only if $\partial^2 c_{ij} / \partial (z_{ij}^k)^2 > 0$, so there must be increasing marginal costs of improving quality. In that case, either an increase in the valuation of quality α^k or an increase in the specific transport costs to the destination market T_i^k will *raise* quality z_{ij}^k . This occurs in particular with an increase in T_i^k due to greater distance, which is related to the well-known ‘‘Washington apples’’ effect.¹¹

Making use of the Cobb-Douglas production function for quality in (2) and the cost function in (3), the second-order conditions are satisfied when $0 < \theta < 1$ which we have already assumed. Further assuming that $0 < \alpha^k \theta < 1$, the first-order condition (5) is readily solved for quality as:

¹¹ The ‘‘Washington apples’’ effect from Alchian and Allen (1964) states that the *relative* price of a higher quality product will fall as a specific transport cost is increased. That effect does not occur in our model because, as noted below in (7), the nominal prices charged by firms of differing productivity and quality to a given destination market are identical. But an increase in the specific transport cost will still lead all firms to increase their quality.

$$\ln z_{ij}^k = \theta \left[\ln T_i^k - \ln(w_i / \varphi_{ij}) + \ln(\alpha^k \theta / (1 - \alpha^k \theta)) \right]. \quad (6)$$

Conveniently, the Cobb-Douglas production function and specific trade costs give us a log-linear form for the optimal quality choice. Since we are allowing $\alpha^k = h(U^k)$ to depend on the utility of the destination market, it follows that richer countries (with higher utility) may import higher quality, as found empirically by Hallak (2006). In addition, quality in (6) is rising in the productivity of the exporting firm, confirming the finding of Schott (2004) that richer (more productive) countries export higher quality goods.¹² Substituting (6) into the cost function (3), we obtain $c_{ij}(z_{ij}^k, w_i) = [\alpha^k \theta / (1 - \alpha^k \theta)] T_i^k$. Thus, the marginal costs of production are proportional to the specific trade costs, which we use below.

Applying the CES expenditure function in (1a) and solving (4) for the optimal choice of the f.o.b. price yields the familiar markup,

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

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

$$p_{ij}^{*k} = T_i^k \left[\left(\frac{1}{1 - \alpha^k \theta} \right) \left(\frac{\sigma}{\sigma - 1} \right) - 1 \right] \equiv \overline{p_i^{*k}}, \quad (7a)$$

$$p_{ij}^k = \tau_i^k T_i^k \left[\left(\frac{1}{1 - \alpha^k \theta} \right) \left(\frac{\sigma}{\sigma - 1} \right) \right] \equiv \overline{p_i^k}. \quad (7b)$$

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

¹² We could write $T_i^k = w_i d_i^k$, where d_i^k is in units of the aggregate factor and depends on distance. In that case, wages w_i (which also depend on productivity) cancel out from (6).

specific transport costs to each market, but are independent of the productivity of the firm j , as indicated by the notation $\overline{p_i^{*k}}$ and $\overline{p_i^k}$. This result is obtained because more efficient firms sell higher quality goods, leading to constant prices in each destination market, and is a razor-edge case between having the largest firms charge low prices (due to high productivity) or high prices (due to high quality). While this razor-edge case simplifies our analytical results, it is not essential to our analysis because we ultimately rely on industry rather than firm-level prices.

We can generalize the cost function in Assumption 2 to take the form $z_{ij}^k = (\varphi_{ij} l_{ij}^k + \psi_{ij})^\theta$, where ψ_{ij} can be interpreted as either plant capability or the factor requirement of another input, as explained in Appendix A. In that case, we no longer find that the prices of firms are constant in a particular destination market, but can be rising or falling in firm productivity. Much of our theoretical analysis goes through in that case, and in particular the log-linear solution for quality as in (6), except that in place of the specific transport cost T_i^k appearing in (6) – which is tightly related to the f.o.b. price from (7a) – we instead have the f.o.b. price plus specific transport cost, $p_{ij}^{*k} + T_i^k$, appearing in (6). In practice it would be difficult to measure this hybrid variable lying in-between the f.o.b. and c.i.f. prices (since the latter also include *ad valorem* trade costs),¹³ so for this reason we do not use the more general cost function.

Combining (6) and (7a) reveals that log quality is a fraction of the log f.o.b. price:

$$\ln z_{ij}^k = \theta \left[\ln(\kappa_1^k \overline{p_i^{*k}}) - \ln(w_i / \varphi_{ij}) \right], \text{ with } \kappa_1^k \equiv \left[\frac{\alpha^k \theta (\sigma - 1)}{1 + \alpha^k \theta (\sigma - 1)} \right]. \quad (8)$$

Thus, to isolate quality from the f.o.b. price we need to know the key parameter θ from the production function for quality, which we estimate in section 4, and productivity-adjusted input

¹³ Alfonso, Moxnes and Opmomolla (2011) provide a method for estimating specific trade costs that relies on firm-level data, which we do not have for our broad sample of goods and countries.

prices, to which we now turn.

3. Solving for Wages and Quality-Adjusted Prices

It would be a formidable challenge to assemble the data on wages, other input prices and firms' productivities needed to directly measure quality in (8) across many goods and countries. In our trade data we will not have such firm-level information. Accordingly, we rely instead on the zero-cutoff-profit condition of Melitz (2003) to solve for the productivity-adjusted wage of the marginal exporter to each destination market, and will thereby obtain quality and quality-adjusted prices.

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 (8), $\hat{P}_i^k \equiv \overline{p}_i^k / [z_i^k (\hat{\varphi}_i^k)]^{\alpha^k}$ denotes the quality-adjusted price for the marginal exporter:

$$\hat{P}_i^k = \overline{p}_i^k \left[(w_i / \hat{\varphi}_i^k) / \kappa_1^k \overline{p}_i^{*k} \right]^{\alpha^k \theta}. \quad (9)$$

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

$$\frac{\hat{X}_i^k}{\text{tar}_i^k \sigma} = f_i^k(\hat{\varphi}_i^k). \quad (10)$$

The term one plus the *ad valorem* tariff tar_i^k appears in the denominator on the left because tariffs must be deducted from revenue before computing profits. Equivalently, we can move the term tar_i^k to the right where it will multiply fixed costs $f_i^k(\hat{\varphi}_i^k)$, which from **A6** are assumed to depend on the cutoff productivity for reasons that we now explain.

The ZCP condition potentially imposes a tight connection between the quality-adjusted prices of two countries i and j selling to the same destination market k . Dividing (10) for these two countries and using the CES demand system,

$$\frac{\hat{X}_i^k}{\hat{X}_j^k} = \left(\frac{\hat{P}_i^k}{\hat{P}_j^k} \right)^{-(\sigma-1)} = \frac{\text{tar}_i^k f_i^k}{\text{tar}_j^k f_j^k}. \quad (10')$$

Thus, if market k has the same import tariffs on countries i and j , and if their fixed costs of exporting are the same, $f_i^k = f_j^k = f^k$, then the export revenue and quality-adjusted prices of the marginal firms from both source countries *are equal*. With a Pareto distribution for productivity, this equality will also apply to the *average* quality-adjusted prices from both source countries to market k .¹⁴ So in that case, the entire difference in observed unit-values between exporters would be attributed to quality.

To avoid this automatic outcome, we shall adopt a more flexible specification for fixed costs. For the firm with productivity $\hat{\varphi}_i^k$, the fixed cost of exporting from country i to k is assumed to be:

$$f_i^k(\hat{\varphi}_i^k) = \left(\frac{w_i}{\hat{\varphi}_i^k} \right) \left(\frac{Y^k}{p^k} \right)^{\beta_0} e^{\beta' F_i^k}, \quad \beta_0 > 0. \quad (11)$$

There are three features of these fixed costs that deserve attention. First, we have written wages on the right of (11) as adjusted for productivity of the ZCP exporter. That is, we are assuming that an exporting firm's productivity applies equally well to variable and fixed costs, as also assumed by Bilbiie, Ghironi, and Melitz (2012) – though in their case, productivity is equal across firms. This specification implies that more productive (marginal) exporters have lower fixed costs and therefore lower quality-adjusted prices from (10'), implying higher quality.

¹⁴ As shown in Appendix B, with a Pareto distribution for firm productivities the average quality-adjusted price to a market is proportional to the quality-adjusted price of the marginal exporter from each country.

The second important feature of the fixed costs in (11) is that we allow them to depend on real expenditure (Y^k/p^k) in the destination market k .¹⁵ This specification follows from the hypothesis of Arkolakis (2010) that small markets have lower fixed costs because it is easier to reach all customers. By adopting this specification, we will prevent very small markets from automatically having the highest import quality because only the most efficient firms can export there. We shall rely on estimates from Arkolakis and others for the parameter β_0 .

The final term appearing in (11) is the exponential of a vector of bilateral variables F_i^k that influence fixed costs, times their coefficients β . In principle these could be any variables that determine the fixed cost of exporting to a market. We will rely on several measures of language similarity between any two countries to measure these, as discussed in Appendix C.

Having specified the fixed costs of exporting, the next step is to use (10)-(11) to solve for productivity-adjusted wages, and substitute that solution into (9) to obtain quality-adjusted prices. To illustrate this solution, let us assume for the moment that firms are homogeneous in their productivities, so that $\hat{\varphi}_i^k$ does not depend on k and denotes the productivity of every firm in country i . This assumption is just an expositional device, and in fact, the solution for quality-adjusted prices is nearly the same once we allow for heterogeneous firms with a Pareto distribution for productivities.¹⁶ We will indicate in the text precisely how the solution changes when we allow for heterogeneous firms, and provide the derivations in that more complex case in Appendix B.

¹⁵ For p^k we use an import unit-value for that good in country k , not adjusted for quality.

¹⁶ As shown by Demidova and Krishna (2007) and Melitz and Redding (2013), with homogeneous firms and fixed costs of exporting, either all firms find it profitable to export or no firms export. Only in a razor-edge case will the ZCP condition apply so that firms are indifferent between exporting or not. Because we are relying on the ZCP condition in our discussion of the homogeneous firms case, we view this discussion as an expositional device only, and show in Appendix B that a very similar solution is obtained with heterogeneous firms.

With the assumption of firm homogeneity, the total exports from country i to k are $X_i^k = \hat{X}_i^k N_i$, where N_i denotes the number of firms in country i . Combining this equation with (9)-(11), we readily obtain the quality-adjusted price,

$$\hat{P}_i^k = \left(\overline{P}_i^k / \left(\kappa_1^k \overline{P}_i^{*k} \right)^{\alpha^k \theta} \right) \left(\frac{X_i^k}{\sigma \text{tar}_i^k N_i} \left(\frac{Y^k}{P^k} \right)^{-\beta_0} e^{-\beta' F_i^k} \right)^{\alpha^k \theta}. \quad (12)$$

This solution for the quality-adjusted price comes from the supply-side of the model, i.e. from the zero-cutoff profit condition. Notice that given the number of firms, exports X_i^k are *positively* related to the quality-adjusted price, in contrast to the demand-side intuition discussed in section 1. That positive relation occurs because when comparing exports from two countries to the same destination market, higher exports per firm are associated with higher fixed costs of exporting, from (10), and therefore with higher productivity-adjusted wages in (11). Hence, quality is lower in (8) and the quality-adjusted price is higher.

A very similar supply-side relation and intuition continues to hold when we allow for heterogeneous firms. In that case, we assume that productivity is Pareto distributed according to **A4** in section 2. With heterogeneous firms, we first integrate the quality-adjusted prices over all firms exporting to country k with productivity greater than $\hat{\varphi}_i^k$. Letting M_i denote the mass of firms in country i , only $M_i[1 - G(\hat{\varphi}_i^k)]$ actually export to country k . Then using the zero-cutoff-profit condition, we show in Appendix B that the average quality-adjusted price \overline{P}_i^k for exports from country i to k is:

$$\overline{P}_i^k = \left(\overline{P}_i^k / \left(\kappa_1^k \overline{P}_i^{*k} \right)^{\alpha^k \theta} \right) \left(\frac{X_i^k / \kappa_2^k \text{tar}_i^k}{M_i (\varphi_i / w_i)^\gamma} \left(\frac{Y^k}{P^k} \right)^{-\beta_0} e^{-\beta' F_i^k} \right)^{\frac{\alpha^k \theta}{(1+\gamma)}} \left(\kappa_2^k \right)^{\frac{1}{1-\sigma}}. \quad (13)$$

with,

$$\kappa_2^k \equiv \frac{\gamma}{[\gamma - \alpha^k \theta (\sigma - 1)]} > 1.$$

Comparing (12) and (13), we see that there are three differences: (i) σ in (12) is replaced by $\kappa_2^k \equiv \sigma\gamma / [\gamma - \alpha^k\theta(\sigma - 1)]$, which includes additional terms that arise from integrating with the Pareto distribution; (ii) N_i in (12) is replaced by $M_i(\varphi_i / w_i)^\gamma$ in (13), which includes a term reflecting the lower-bound of productivity relative to country wages;¹⁷ (iii) the final exponent $\alpha^k\theta$ in (12) is replaced by $\alpha^k\theta / (1 + \gamma)$ in (13), which includes the Pareto parameter γ . This third change arises because only a subset of firms $M_i[1 - G(\hat{\varphi}_i^k)]$ actually export from country i to k , and because this set of firms is endogenous, it introduces an additional *extensive* margin of substitution in trade between them that is governed by the Pareto parameter γ .

With heterogeneous firms, we see once again in (13) that an increase in exports to a market, given the mass of firms, raises the relative quality-adjusted price. That occurs because an increase in relative exports means that less-efficient firms are exporting to that market, and therefore average quality falls. Again, that relationship sounds contrary to the demand-side intuition discussed in section 1: given nominal prices, higher sales to a market should mean higher quality. In fact, that intuition continues to hold in our model, and we shall use it below in conjunction with (13) to solve for the quality-adjusted prices.

Quality-Adjusted Export Prices

Let us return to the expositional assumption that firms are homogeneous. Then in the zero-cutoff-profit condition (10), the firm-level sales \hat{X}_i^k are obtained from total exports as $\hat{X}_i^k = X_i^k / N_i$, which in turn equals CES demand from the expenditure function in (1):

¹⁷ With country wages following the lower bound of productivity in equilibrium, this extra term should not be too important. We control for it by including the labor force in our empirical specification; see (21).

$$\hat{X}_i^k = \frac{X_i^k}{N_i} = \left(\frac{\hat{P}_i^k}{P^k} \right)^{-(\sigma-1)} Y^k, \quad (14)$$

where P^k is the price index corresponding to the CES expenditure function in (1). Consider dividing (14) for two countries i and j selling to the same market k , to solve for the relative quality-adjusted export prices,

$$\frac{\hat{P}_i^k}{\hat{P}_j^k} = \left(\frac{X_i^k / N_i}{X_j^k / N_j} \right)^{\frac{-1}{\sigma-1}}. \quad (15)$$

Given an empirical specification of the number of products available from each country, and the elasticity of substitution, we could use (15) to determine the relative quality-adjusted export prices to each market. This equation embodies the demand-side intuition that goods with higher market shares are assigned higher quality and hence lower quality-adjusted price, as used by Khandelwal (2010) and Hallak and Schott (2011).

Our framework with zero profits for the marginal exporter allows for a tighter solution for the quality-adjusted export prices, however. We can substitute the demand-side equation (14) into the supply-side equation (12) to eliminate exports X_i^k , in which case the number of products N_i cancels out and we readily solve for the ratio:

$$\frac{\hat{P}_i^k}{\hat{P}_j^k} = \left(\frac{\overline{p}_i^k / \left(\overline{tar}_i^k \overline{p}_i^{*k} e^{\beta' F_i^k} \right)^{\alpha^k \theta}}{\overline{p}_j^k / \left(\overline{tar}_j^k \overline{p}_j^{*k} e^{\beta' F_j^k} \right)^{\alpha^k \theta}} \right)^{\frac{1}{1+\alpha^k \theta (\sigma-1)}}. \quad (16)$$

Comparing (15) with (16), it is apparent that we obtain a different solution for quality-adjusted export prices when the supply-side of the model is also used: in (16), the quality-adjusted prices are tightly pinned down by the c.i.f. and f.o.b. prices that appear on the right, as well as by tariffs

and the fixed cost terms. Remarkably, the relative number of products N_i/N_j does not enter (16), which occurs because the ZCP condition is solving for the *per-firm* exports $\hat{X}_i^k / \hat{X}_j^k$, which also appears in the demand equation (14), and so these supply and demand conditions together are eliminating the unobserved number of firms. Eliminating this variable is the key simplification that we obtain by using the supply side of our model.¹⁸

When we allow for heterogeneous firms with a Pareto distribution for productivities, the solution for quality-adjusted export prices is the same as in (16). As shown in Appendix B, the demand equation (14) can be re-expressed in a form that is close to a gravity equation:

$$\frac{X_i^k}{M_i(\varphi_i / w_i)^\gamma} = \left(\frac{\overline{P_i^k}}{P^k} \right)^{-(\sigma-1)(1+\gamma)} \left(Y^k \right)^{(1+\gamma)} \left(\sigma \kappa_2^k \text{tar}_i^k \left(\frac{Y^k}{P^k} \right)^{\beta_0} e^{\beta' F_i^k} \right)^{-\gamma}, \quad (17)$$

where $\overline{P_i^k}$ is the average quality-adjusted price. 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. Exports are divided by the mass of potential exporters M_i on the left, analogous to dividing by N_i in (14), even though only a fraction of firms $M_i[1 - G(\hat{\varphi}_i^k)]$ actually export from i to k . That extensive margin of substitution is reflected in the exponent $-(\sigma - 1)(1 + \gamma)$ which appears on the relative price in (17): we will refer to this term as the “elasticity of trade”, and comparing (14) with (17), we see that this elasticity is higher in absolute value when the extensive margin is taken into account.

¹⁸ Of course, if the number of firms takes on their equilibrium values, then (15) and (16) would give the same solution for the relative quality-adjusted export price. The problem in practice is that it is very difficult to have a parsimonious specification for the number of firms that gives a similar solution in (15) and (16), as we shall demonstrate in section 5.

Continuing with heterogeneous firms case, we can substitute the demand-side equation (17) into the supply-side equation (13) to eliminate exports X_i^k , in which case the mass of firms M_i again cancels out. Taking the ratio of relative quality-adjusted prices $\overline{P_i^k} / \overline{P_j^k}$ we obtain exactly the same expression as (16), which now applies to the *average* quality-adjusted prices, i.e. integrating over all firms with productivities above the ZCP exporter:

$$\frac{\overline{P_i^k}}{\overline{P_j^k}} = \left(\frac{\overline{p_i^k} / \left(\overline{tar_i^k} \overline{p_i^{*k}} e^{\beta' F_i^k} \right)^{\alpha^k \theta}}{\overline{p_j^k} / \left(\overline{tar_j^k} \overline{p_j^{*k}} e^{\beta' F_j^k} \right)^{\alpha^k \theta}} \right)^{\frac{1}{1 + \alpha^k \theta (\sigma - 1)}}. \quad (18)$$

We will use this ratio to measure the relative quality-adjusted export prices of countries i and j selling to each market k . This relative price is similar in spirit to Khandelwal (2010) and Hallak and Schott (2011), who measure export prices to the United States. We shall repeat this for each destination market k , and then aggregate over destinations and over goods, as discussed in Appendix D. The key message from this section is that when measuring quality-adjusted *export* prices, we can go beyond the pure demand-side measurement in (15) by also using the ZCP condition on the supply-side, thereby obtaining the tight solution in (18).

Quality-Adjusted Import Prices

We shall also want to measure quality-adjusted import prices, which has not been done before in the literature. In that case, we consider each source county i selling to two destination markets k and l , and form the ratio $\overline{P_i^k} / \overline{P_i^l}$, which measures the quality-adjusted import price for country k relative to l . We rely on the supply-side equation (13) to obtain the ratio $\overline{P_i^k} / \overline{P_i^l}$, and we find once again that the mass of exporters M_i cancels out. We still find, however, that the ratio of (13) involves two different taste parameters α^k and α^l , reflecting the differing weights

that destination markets k and l put on quality. We do not want our measurement of quality-adjusted prices to depend on differing preferences across countries, so we replace the taste parameters α^k and α^l with the average value $\bar{\alpha}$ for all countries importing the good.¹⁹ Then adding the goods and time subscripts g and t , we measure the ratio of (13) for a country i selling to two destinations k and l as:

$$\frac{\overline{P_i^k}}{\overline{P_i^l}} = \frac{\left(\frac{\overline{p_i^k}}{\overline{p_i^l}} / \left(\frac{\kappa_1^k \overline{p_i^{*k}}}{\kappa_1^l \overline{p_i^{*l}}} \right)^{\bar{\alpha}\theta} \right)^{\frac{\bar{\alpha}\theta}{(1+\gamma)}} \left(\frac{X_i^k / \kappa_2^k \text{tar}_i^k \left(\frac{Y^k}{P^k} \right)^{\beta_0} e^{\beta' F_i^k}}{X_i^l / \kappa_2^l \text{tar}_i^l \left(\frac{Y^l}{P^l} \right)^{\beta_0} e^{\beta' F_i^l}} \right)^{\frac{\bar{\alpha}\theta}{(1+\gamma)}}}{\left(\frac{\kappa_2^k}{\kappa_2^l} \right)^{\frac{1}{1-\sigma}}} . \quad (19)$$

Comparing the relative export price in (18) with the relative import price in (19), it is apparent that the export prices in (18) have the smaller exponent $1/[1 + \alpha^k \theta(\sigma - 1)] < 1$ on the ratio of c.i.f. to f.o.b. prices. In our estimates, this exponent has a median value less than 0.25 and over 98 percent of estimates across industries and countries are less than 0.5. This is one reason that we shall find that the quality-adjusted export prices differ by less than the quality-adjusted import prices across countries; another reason is the extra terms appearing on the right of (19), discussed below. The smaller exponent on the c.i.f./f.o.b. ratio of export prices occurs because we find that consumers in a given destination market have a high degree of substitution between the goods from different countries: so in order to have the level of trade consistent with the data, we will find that quality-adjusted export prices cannot differ by that much. But this intuition does not apply to the relative quality-adjusted import prices in (19), which compare

¹⁹ According to Fisher and Shell (1972), with changing preferences (in this case changing between countries), a suitable approach is to compute a geometric mean of price indexes that first uses one country's preferences and then uses the other's. We have also implemented the Fisher-Shell approach for our import indexes, as discussed in Appendix D, and find similar results to using the average preference for quality $\bar{\alpha}$.

country i selling to two destinations k and l . In that case there is no direct consumer substitution between the products, and the quality-adjusted import prices are instead based on the supply relation from (13). It follows that these import prices will have greater dispersion across countries than the relative export prices.

The relative import prices also depend on a number of additional terms besides the c.i.f./f.o.b. price ratio. Most important, the relative import prices depend on destination market expenditure Y^k in two ways. On the one hand, higher expenditure leads to greater exports X_i^k to that country. The marginal exporters will be less efficient, producing lower quality with higher quality-adjusted price. That is the negative supply-side relation between exports and quality that we have already discussed. This effect is offset by higher real expenditure (Y^k/p^k) in (19) leading to higher fixed costs. In that case the marginal exporter must be more efficient, leading to higher quality and lower quality-adjusted price. The strength of these two opposing forces depends on the parameter β_0 . That parameter is estimated with firm-level export data by Arkolakis (2010), Eaton, Kortum and Kramarz (2011) and Eaton, Kortum and Sotelo (2012), who obtain $\beta_0 \approx 0.35$. In Appendix C we discuss how our specification of fixed costs in (11) – depending on the productivity of the cutoff exporter – maps into the same firm-level data, and conclude that β_0 in our model lies between zero and 0.35, depending on the Pareto parameter γ for the good in question. So we use this calibration for β_0 in the calculation of the relative import prices in (19). The estimation of the Pareto parameter, the elasticity of substitution σ , and the quality parameter θ are discussed in the next section.

4. Data and Estimation

Data

Our primary dataset is the United Nations' Comtrade database, used to obtain export and import data for 185 countries from 1984-2011. We compute the bilateral f.o.b. unit values of traded goods using reports from the exporting country. By focusing on the exporters' reports we ensure that these unit values are calculated prior to the inclusion of any costs of shipping the product. The bilateral c.i.f. unit values are calculated similarly using importers' trade reports. Since these unit values include the costs of shipping, we need only add the tariff on the good to produce a tariff-inclusive c.i.f. unit value. To do this we obtain the *ad valorem* tariffs associated with Most Favored Nation status or any preferential status from raw TRAINS data and from the World Trade Organization's (WTO) Integrated Data Base (IDB), which we have expanded upon using tariff schedules from the *International Customs Journal* and the texts of preferential trade agreements obtained from the WTO's website and other online sources. We provide further details in Appendix C.

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

of trade was less than \$50,000 in constant 2005 dollars.

More generally, to reconcile the wide variation in the observed unit values with our model, we assume that the f.o.b. and duty-free c.i.f. unit values, denoted by uv_{igt}^{*k} and uv_{igt}^k with goods subscript g and time subscript t , are related to the f.o.b. and tariff-inclusive c.i.f. prices by:

$$\ln uv_{igt}^{*k} = \ln \overline{p_{igt}^{*k}} + u_{igt}^{*k} \quad \text{and} \quad \ln uv_{igt}^k = \ln(\overline{p_{igt}^k} / \overline{tar_{igt}^k}) + u_{igt}^k, \quad (20)$$

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

Estimation

We adapt Feenstra's (1994) GMM method to estimate the parameters of the model. To achieve this we take the ratio of the demand equation (17) for two countries i and j selling to destination k , and substitute for the relative quality-adjusted export prices in (19), while adding subscripts for goods g and time t . Because the demand equation contains the unobserved mass of potential exporters, we need to control for this mass. We estimate the labor force L_{igt} employed in producing exports of good g in country i as country i population multiplied by country i exports of good g divided by country i GDP. We then model the mass of potential exporters as depending on L_{igt} and country fixed effects:

$$\ln[M_{igt}(\varphi_{igt}/w_{igt})^\gamma] = \delta_{0g} \ln L_{igt} + \delta_{ig} + \varepsilon_{igt}^k, \quad (21)$$

where ε_{igt}^k is a random error. We also use (20) to replace the c.i.f. and f.o.b. prices with their respective unit values. Then from (17) and (19)–(21), we obtain the difference between exports from countries i and j selling to destination k :

$$\begin{aligned} \ln X_{igt}^k - \ln X_{jgt}^k = & -A_g^k \left[\left(\ln(\text{tar}_{igt}^k uv_{igt}^k) - \ln(\text{tar}_{jgt}^k uv_{jgt}^k) \right) - \alpha_g^k \theta_g \left(\ln uv_{igt}^{*k} - \ln uv_{jgt}^{*k} \right) \right] \\ & + \delta_{0g} (\ln L_{git} - \ln L_{gjt}) + \delta_{ig} - \delta_{jg} - B_g^k [\ln \text{tar}_{igt}^k + \beta_g' (F_i^k - F_j^k)] + \varepsilon_{igt}^k - \varepsilon_{jgt}^k, \end{aligned} \quad (22)$$

where:

$$A_g^k \equiv \frac{(\sigma_g - 1)(1 + \gamma_g)}{1 + \alpha_g^k \theta_g (\sigma_g - 1)}, \quad \text{and} \quad B_g^k \equiv \frac{\gamma_g - \alpha_g^k \theta_g (\sigma_g - 1)}{1 + \alpha_g^k \theta_g (\sigma_g - 1)}. \quad (23)$$

We add a simple supply specification in Appendix E, whereby the specific and iceberg trade costs depend on distance and the quantity traded, and iceberg trade costs also depend on *ad valorem* tariffs. Feenstra (1994) assumed that the supply shocks and demand shocks are uncorrelated. That assumption seems unlikely to hold with unobserved quality, since a change in quality could shift both supply and demand. But here, the demand errors and the supply errors are the residuals after *taking into account* quality. So the assumption that they are uncorrelated seems much more acceptable, and is the basis for the GMM estimation.

Two features of the estimating equation (22) deserve attention. First, notice that the c.i.f. unit values appear with the negative coefficient $-A_g^k$ in this gravity equation, whereas the f.o.b. unit values appear with a *positive* coefficient $A_g^k \alpha_g^k \theta_g$. The f.o.b. unit values reflect product quality in the equation, and conditional on the c.i.f. unit value, higher quality leads to higher demand, which explains why the f.o.b. coefficient is positive. The key to successful estimation will be to obtain this sign pattern on the unit values.

Second, not all the parameters are identified without additional information. In particular, we estimate $B_g^k \beta_g'$ in (22) but not these coefficients alone. If we do not identify B_g^k , then we cannot solve for σ_g and γ_g . We resolve this issue as in Chaney (2008), by using estimates of $\zeta_g^{US} = \gamma_g / [\alpha_g^{US} \theta_g (\sigma_g - 1)]$ from regressions of firm rank on size for each SITC sector in the U.S., where we further normalize $\alpha_g^{US} \equiv 1$.²⁰ Then for other countries, $\zeta_g^k = \gamma_g / [\alpha_g^k \theta_g (\sigma_g - 1)]$
 $\Leftrightarrow \zeta_g^k \alpha_g^k = \gamma_g / [\theta_g (\sigma_g - 1)] = \zeta_g^{US} \alpha_g^{US} = \zeta_g^{US}$. It follows that γ_g is obtained as $\zeta_g^{US} \theta_g (\sigma_g - 1)$.

A final parameter that is difficult to identify without additional information is α_g^k , which is the preference for quality in the expenditure function (1). But conveniently, this parameter can be estimated from simple price regressions, estimated in Appendix E. From (7a), the f.o.b. price is increasing in the destination country's preference for quality α_g^k , which we can model as an increasing function of the destination country's per-capita real income with coefficient λ_g . It is well known from Hallak (2006) that the unit-value of imports is positively related to a country's per-capita income, which identifies α_g^k . These price regressions depend on having preliminary estimates of σ_g and θ_g that come from estimating (22) when all countries have the same preference for variety, $\alpha_g^k \equiv 1$. Using these preliminary estimates of σ_g and θ_g , we then estimate the price regressions to obtain improved values for α_g^k . These improved values of α_g^k are substituted into (23), and we re-estimate (22) to obtain new estimates for σ_g and θ_g . We iterated this procedure several times and found that the distribution of estimates for σ_g and θ_g quickly converged.

²⁰ We thank Thomas Chaney for providing these estimates for 3-digit SITC Rev. 3 sectors for the United States, which we concurred to 3-digit SITC Rev. 2 sectors. In Chaney (2008), this parameter equals $\zeta^k = \gamma / (\sigma - 1)$, and we discuss in Appendix B why it equals $\zeta^k = \gamma / [\alpha^k \theta (\sigma - 1)]$ in our model. The normalization $\alpha_g^{US} \equiv 1$ is harmless because α_g^k always appears multiplied by θ , so $\alpha_g^{US} \equiv 1$ fixes the value for θ in our estimates.

Parameter Estimates

Estimation is performed for each 4-digit SITC Revision 2 good (which we also refer to as an industry) using bilateral trade between all available country pairs during 1984-2011. There are 12.5 million observations with data on both the c.i.f. and f.o.b. unit values that passed the data-cleaning criteria detailed above, excluding those goods with fewer than 50 observations. We perform the GMM estimation on 712 industries as shown in the first row of Table 1.²¹ The median estimate of σ_g is 6.07, not counting seven industries with an inadmissible value less than unity; the median estimate of γ_g is 8.43, not counting the same seven industries with an inadmissible value; and the median estimate of θ_g is 0.61, not counting four cases with an inadmissible value less than zero or greater than unity. For inadmissible values or for SITC industries with fewer than 50 observations, we replace the parameter estimates with the median estimate from the same 3-digit or 2-digit SITC industry, after which we find the median estimates shown in the last row of Table 1 for 924 industries.

The frequency distribution of parameter estimates are illustrated in Figures 1-3. Our

Table 1: Median Parameter Estimates

GMM Estimation Method with:	Number of SITC industries	σ	γ	θ
Dropping SITC4 with < 50 observations	712	6.07	8.43	0.61
No. of inadmissible parameters	8	7	7	4
Filling in SITC4 with < 50 observations or inadmissible parameters	924	5.82	7.78	0.61

²¹ In each industry we use only the most common unit of measurement, which is nearly always kilograms.

Figure 1: Frequency Distribution for Estimates of σ_g
(Note: Estimates are right-censored for presentation purposes only)

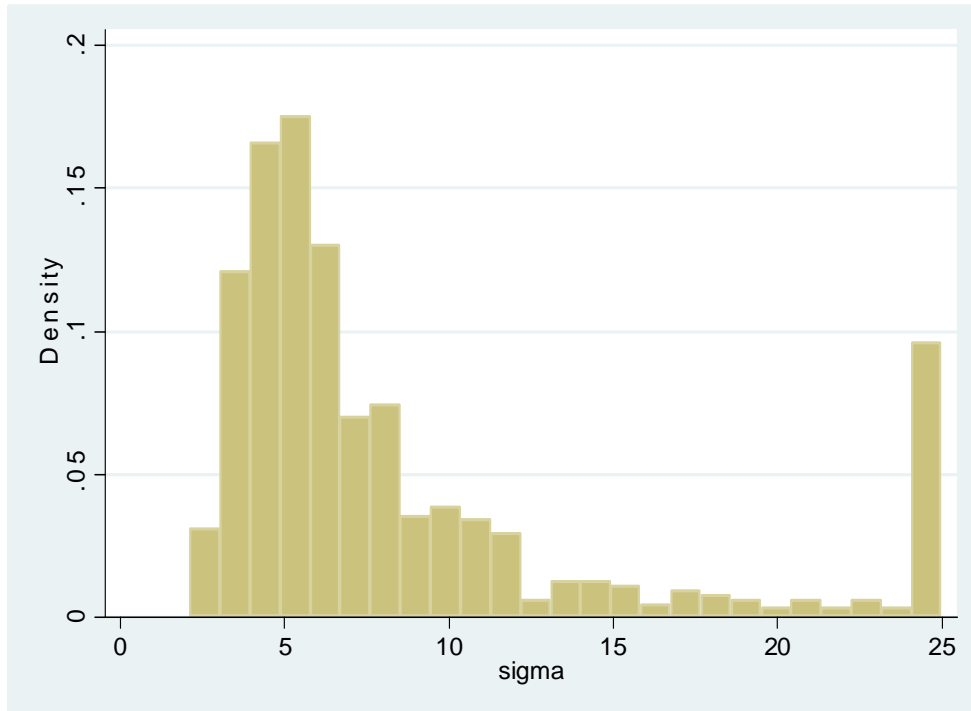


Figure 2: Frequency Distribution for Estimates of γ_g
(Note: Estimates are right-censored for presentation purposes only)

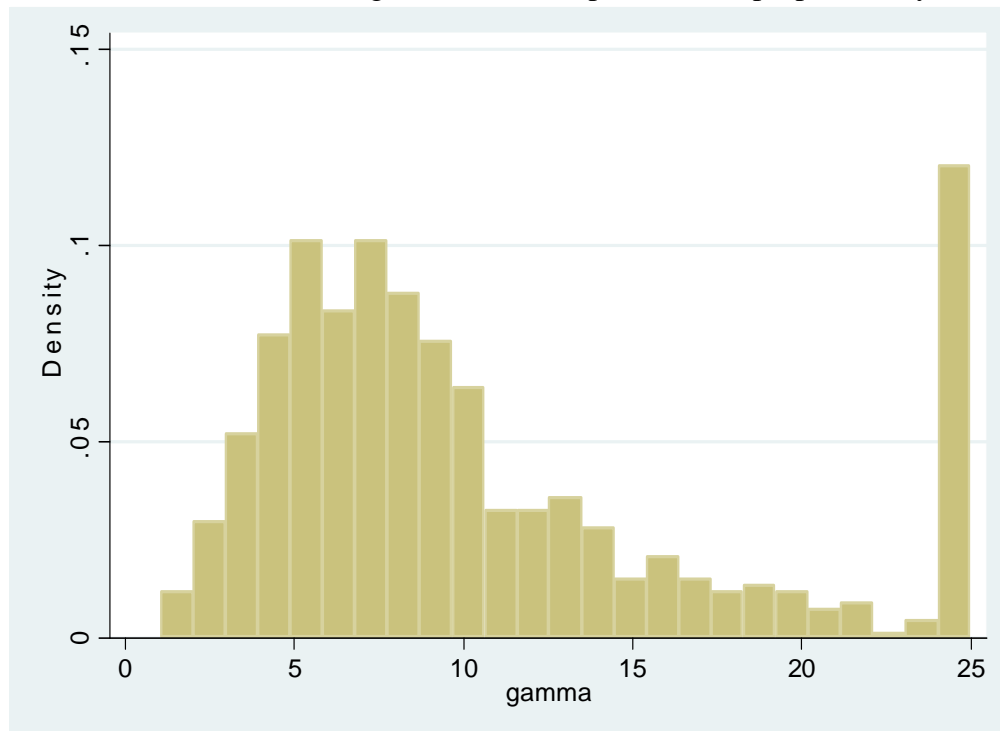
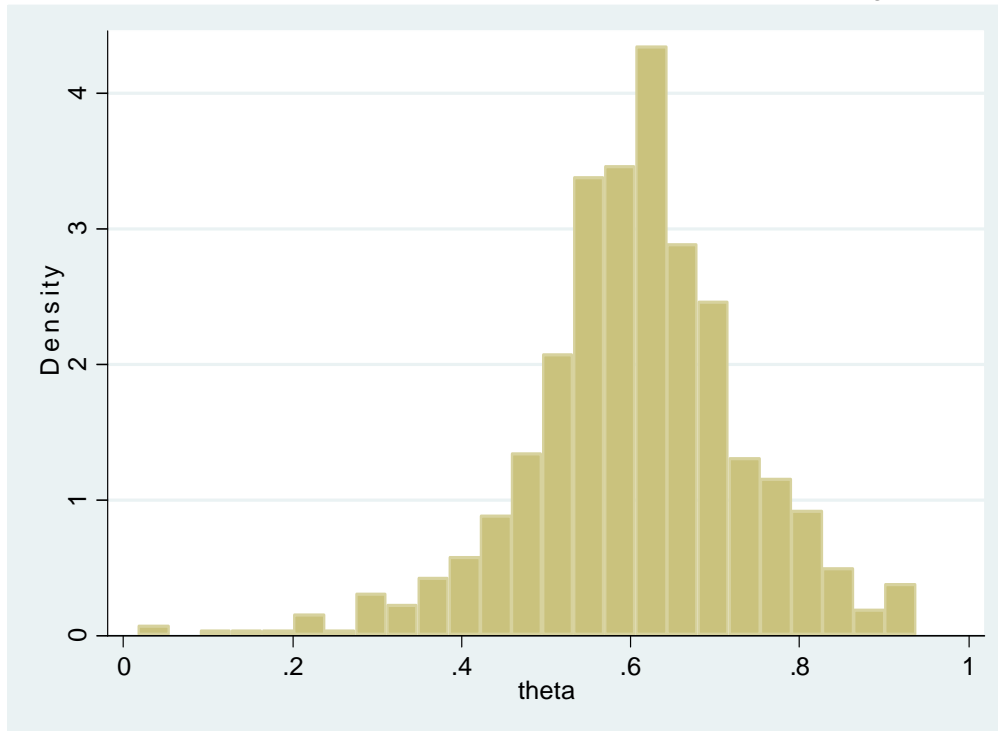


Figure 3: Frequency Distribution for Estimates of θ_g 

median estimate for the elasticity of substitution σ_g is higher than estimated by Broda and Weinstein (2006) for the United States. We have found that our higher value comes from using worldwide trade data and correcting for quality, and from using an empirical specification that is more robust to measurement error since we do not take differences over time and instead include source-country fixed effects in our estimation of (25).²² Our median estimate for the Pareto parameter γ is quite close to that reported by Eaton and Kortum (2002), who also considered bilateral trade between many countries.²³

We know of no other estimate of θ_g . Crozet, Head and Meyer (2012) study firm-level

²² Destination country fixed effects are implicitly included, too, because (20) is specified as the difference between countries i and j exporting to country k .

²³ This median estimate is higher, however, than the recent results of Simonovska and Waugh (2011, 2012).

data for the champagne industry to estimate key parameters of a Melitz (2003) model with quality. They combine export data with expert ratings of the overall quality of each champagne producer on a 1- to 5-star scale. The estimated cost (proportional to f.o.b. price) for 5-star producers is 68 percent higher than for 1-star producers. Though there is no translation of the discrete star-rating to how consumers evaluate the quality of champagne, this estimate appears consistent with a fairly high value of theta – quality increases quite substantially with the use of more or better inputs.

5. Indexes of Quality-Adjusted Price and Quality

The quality-adjusted relative export prices are obtained from (18) and import prices from (19), where we replace the c.i.f. price appearing there by the tariff-inclusive c.i.f. unit value, $uv_{igt}^k tar_{igt}^k$ as in (20), and the f.o.b. price by the f.o.b. unit value uv_{igt}^{*k} . Each of these are then aggregated over partner countries, and from 4-digit SITC to the Broad Economic Categories (BEC), to obtain overall indexes of quality and quality-adjusted prices of exports and imports for each country and year in our dataset. The formula we shall use for aggregation is the so-called GEKS method,²⁴ which is a many-country generalization of Fisher Ideal indexes. We apply a two-stage aggregation procedure over partner countries and then over goods, resulting in an aggregate export and import unit-value for each country relative to the U.S. We refer to the GEKS index of unit values as the “price index” and the GEKS index of quality-adjusted unit values as the “quality-adjusted price index”. Our final step is to divide the former by the latter – for each country, year and BEC– to obtain the index of export or import quality.

²⁴ Named after Gini, Eltetö and Köves, and Szulc. We refer the reader to Balk (2008) and Deaton and Heston (2010) for a modern treatment and details of these historical references. We employ the GEKS procedure here because it is commonly used by statistical agencies, including the ICP and PWT.

Export Prices and Quality

Before showing our results on the export side, we begin by using only the demand side of our model to construct the quality-adjusted prices in (15) for 2007. It is evident that this formula is very sensitive to the specification of the number of exporting firms in each country, or N_i in the homogeneous firms case. We illustrate this by making two different assumptions about N_i : (i) N_i is proportional to countries' population (similar to Khandelwal (2010)); and (ii) N_i is proportional to countries' aggregate non-services value-added. In Figure 4 we show the raw unit-value indexes (top panel) together with export quality indexes when N_i is assumed proportional to population (middle panel) and non-services value-added (bottom panel). In all cases we normalize the world average unit value to unity. The middle panel of Figure 4 reveals quality to be positively correlated with per-capita GDP (correlation coefficient = 0.41), while the bottom panel exhibits virtually no correlation (correlation coefficient = -.03).²⁵

In fact, the sensitivity of quality estimates to our assumptions about N_i may be greater than appears in Figure 4. Excluding small countries (population less than 1 million) that account for the bulk of outliers, these correlations become 0.49 and -0.34 respectively. Without good information, demand-side estimates of quality may largely reflect the researcher's assumptions about the number of firms. Comparing the bottom two panels of Figure 4 with the top panel, it is visually apparent that both demand-side quality estimates vary much more than the unit-value indexes. As a result, the quality-adjusted price indexes in Figure 5 (first using population to proxy the number of exporters, and then non-services value added) show substantial variation across countries: greater than the original unit-value indexes in the top panel of Figure 4.

²⁵ In all figures we exclude St Vincent and the Grenadines, which has very high export prices driven by exports (likely re-exports) of yachts to Greece and Italy and color televisions to Trinidad and Tobago.

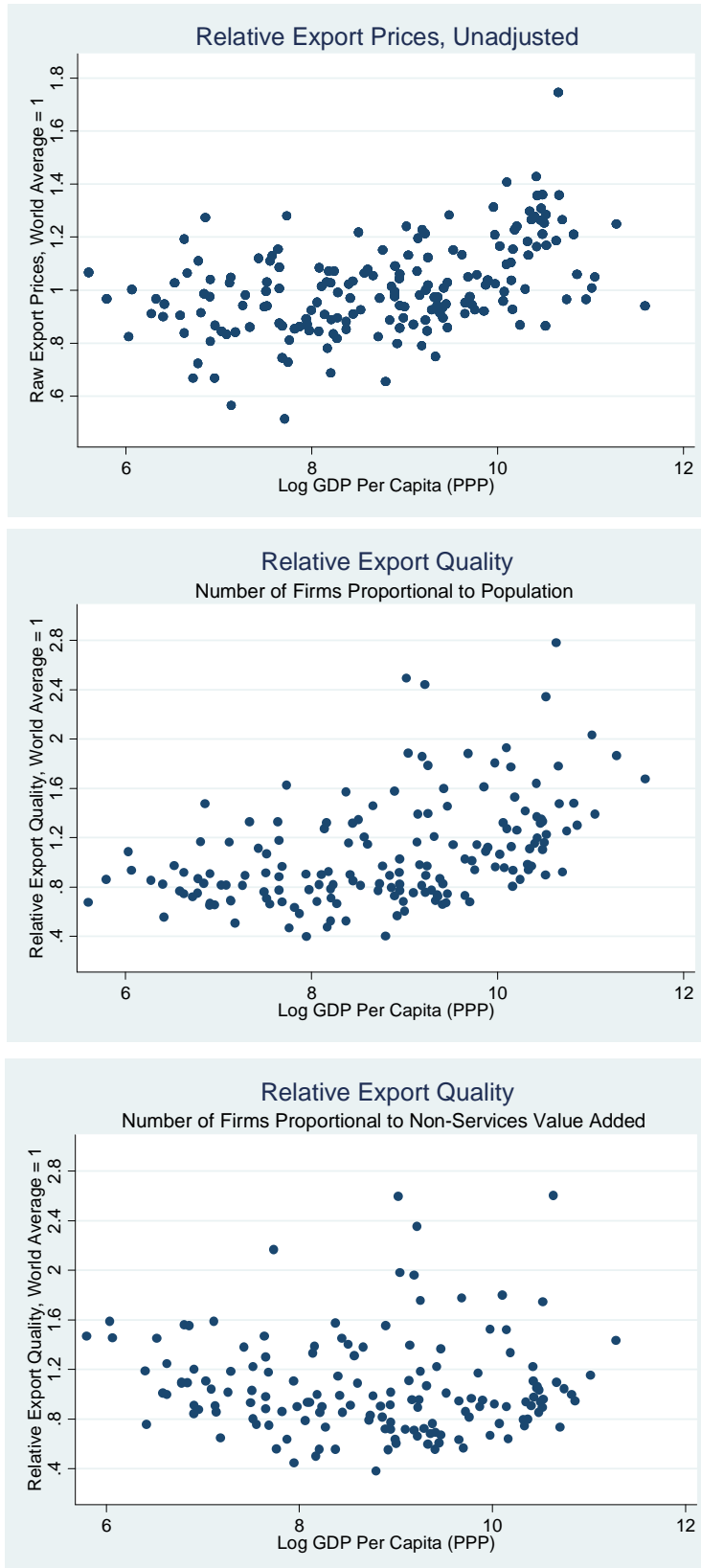
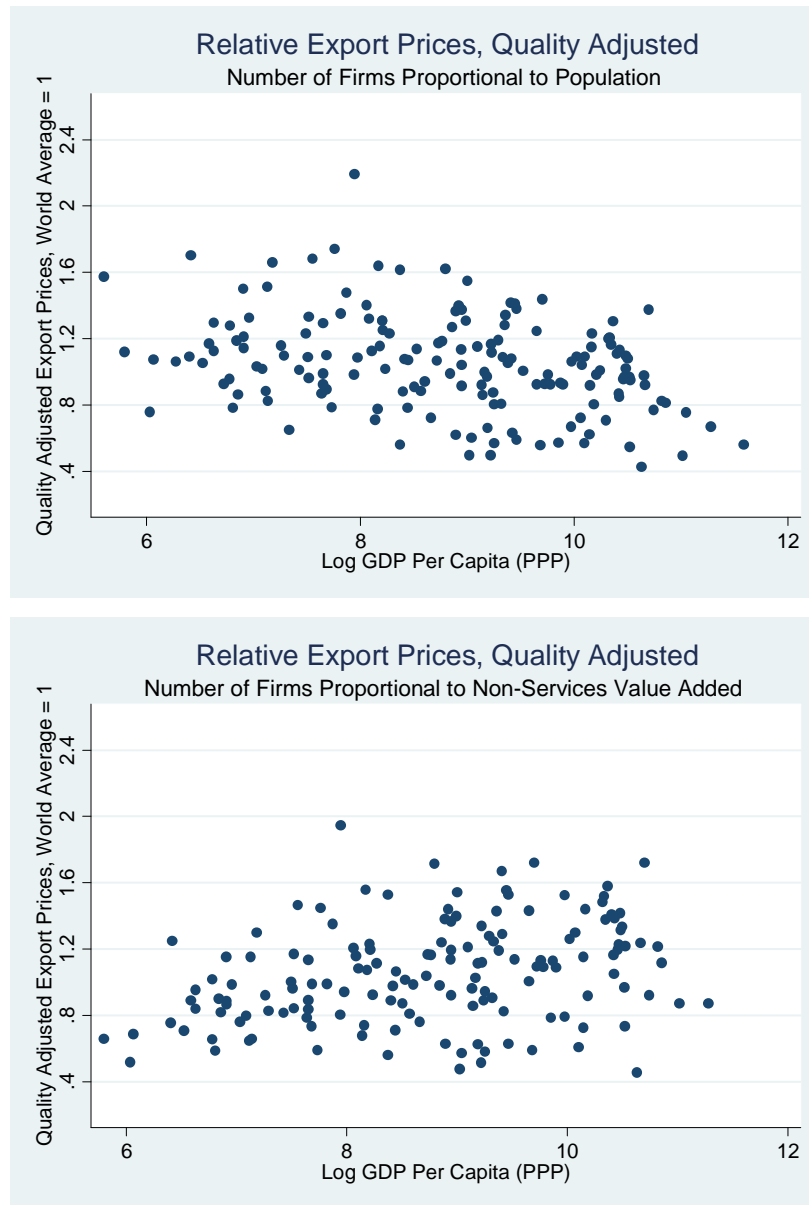
Figure 4: Raw Export Prices and Demand-Side Estimates of Export Quality, 2007

Figure 5: Demand-Side Estimates of Quality-Adjusted Export Prices, 2007

We can contrast these results obtained from the demand side of our model with the quality-adjusted prices in (18), obtained from the demand and supply sides. The quality indexes and the quality-adjusted price indexes for 2007 are shown in Figure 6. Comparing the top panel of Figure 6 with the top panel of Figure 4, it is visually apparent that the quality indexes are now similar to the unit-value indexes, and as a result, the quality-adjusted prices (second panel in

Figure 6: Supply and Demand Based Estimates of Export Quality and Quality-Adjusted Export Prices, 2007

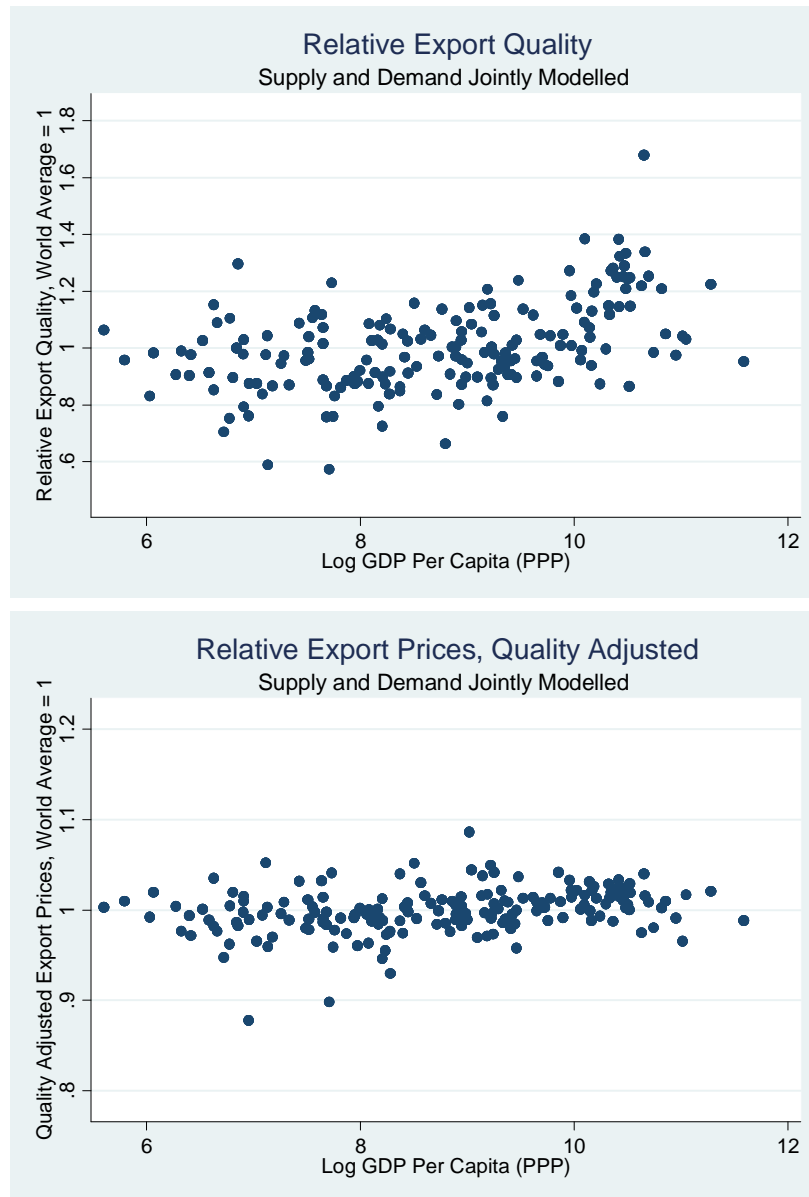


Figure 6) show much less variation than those obtained from the demand side only (in Figure 4).

We offer two reasons for this difference in results. First, the demand-side formula in (15) depends on trade *values* on the right, which can differ by many orders of magnitude for two countries selling to a given destination; in contrast, the c.i.f. and f.o.b. prices appearing on the right of (18) do not differ as much in the data. Second, while this potentially large difference in

trade values can be offset by the estimated number of firms exporting from each country, in practice it is difficult to get reliable estimates of that number, limiting researchers' ability to construct quality-adjusted prices from the demand side alone.²⁶

Turning to other results, we notice that developed countries tend to export more expensive goods (top panel of Figure 4), and we estimate these goods to be of higher than average quality (top panel of Figure 6). The quality adjusted-price (second panel of Figure 6), about which we have less strong priors, tends to be only slightly higher for developed countries, indicating that most of the higher export price for developed countries is explained by quality.

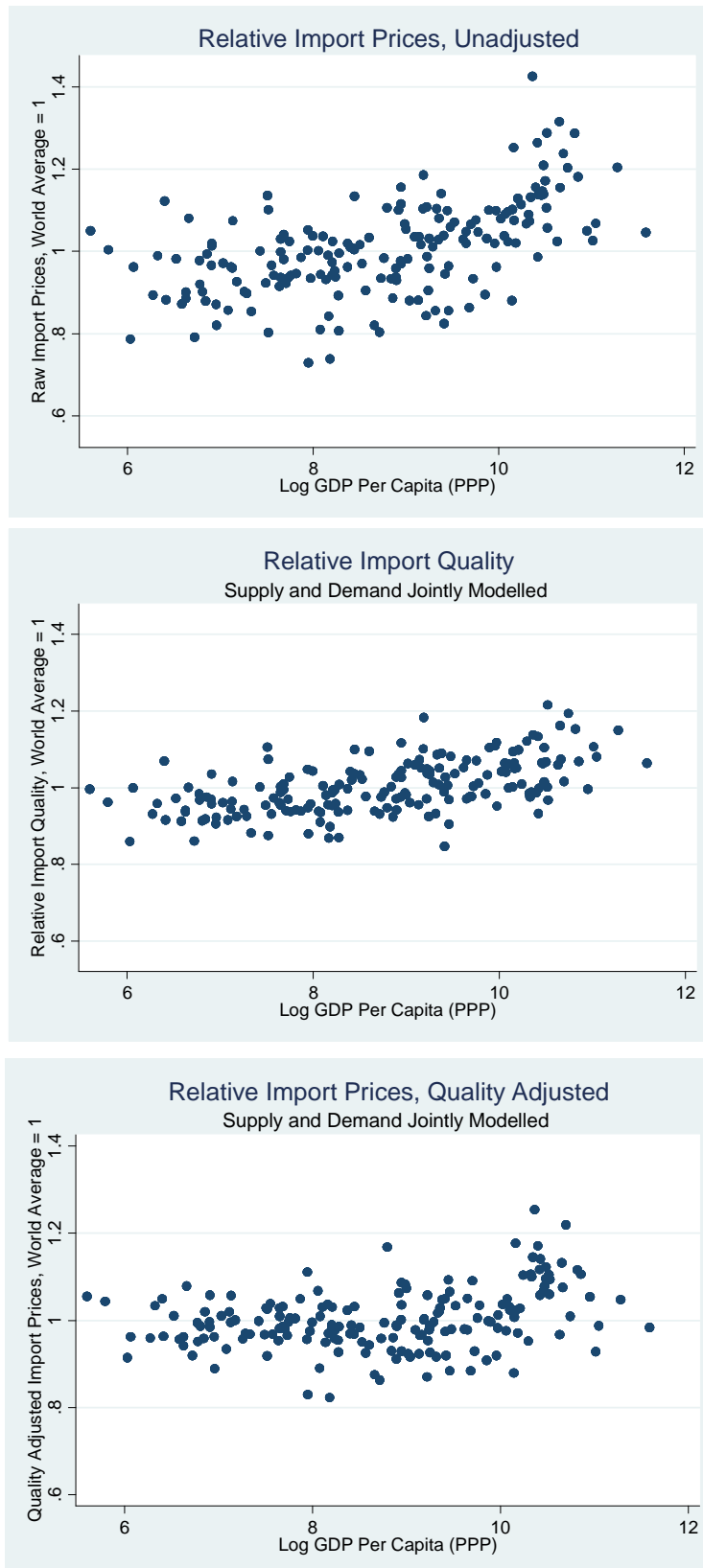
Import Prices and Quality

We illustrate a similar exercise for import prices in Figure 7, but we do not attempt a comparison with the demand-side alone.²⁷ Developed countries import more expensive items (top panel) that are of higher quality (second panel). Quality-adjusted import prices (third panel) increase noticeably with the importing country's GDP per capita. This pattern is due to an interaction of preferences for quality and the rising marginal cost of producing quality. Rich countries tend to prefer higher quality goods, which enter the import quality-adjusted price in (19) via κ_{1g}^k and κ_{2g}^k . But our estimates of θ_g between zero and unity means, from (4), that there is an amplified effect of quality on increasing the marginal cost, so that higher quality induced by a preference for quality leads to a higher quality-adjusted price.

²⁶ As explained in note 18, if we obtained estimates of the number of firms that equaled their equilibrium values, then the quality-adjusted prices obtained from (15) and (18) would be identical.

²⁷ As noted earlier, since Schott (2004), Hallak and Schott (2011) and Khandelwal (2010) all focus on exports to the United States, they do not construct indexes of import prices calculated by comparing prices for a given country selling to two destinations. More generally, it is not possible to go immediately from (14) to a simple specification of quality-adjusted import prices, because the CES price index as well as income of each destination country would enter the formula.

Figure 7: Raw Import Prices and Supply and Demand Based Estimates of Import Quality and Quality-Adjusted Import Prices, 2007



It is evident that the variation in quality-adjusted import prices in Figure 7 is much greater than for export prices in Figure 6. Numerically, this occurs for two reasons. First, as noted above, the c.i.f./f.o.b. ratio of export unit values on the right of (18) has an exponent significantly less than unity, which reflects substitution between suppliers and tends to mute those price differences on the export side; but that does not occur on the import side, where only the f.o.b. price on the right of (19) has an exponent less than unity. Hence, the raw differences in unit values across countries show up more in the quality-adjusted prices for imports than exports.

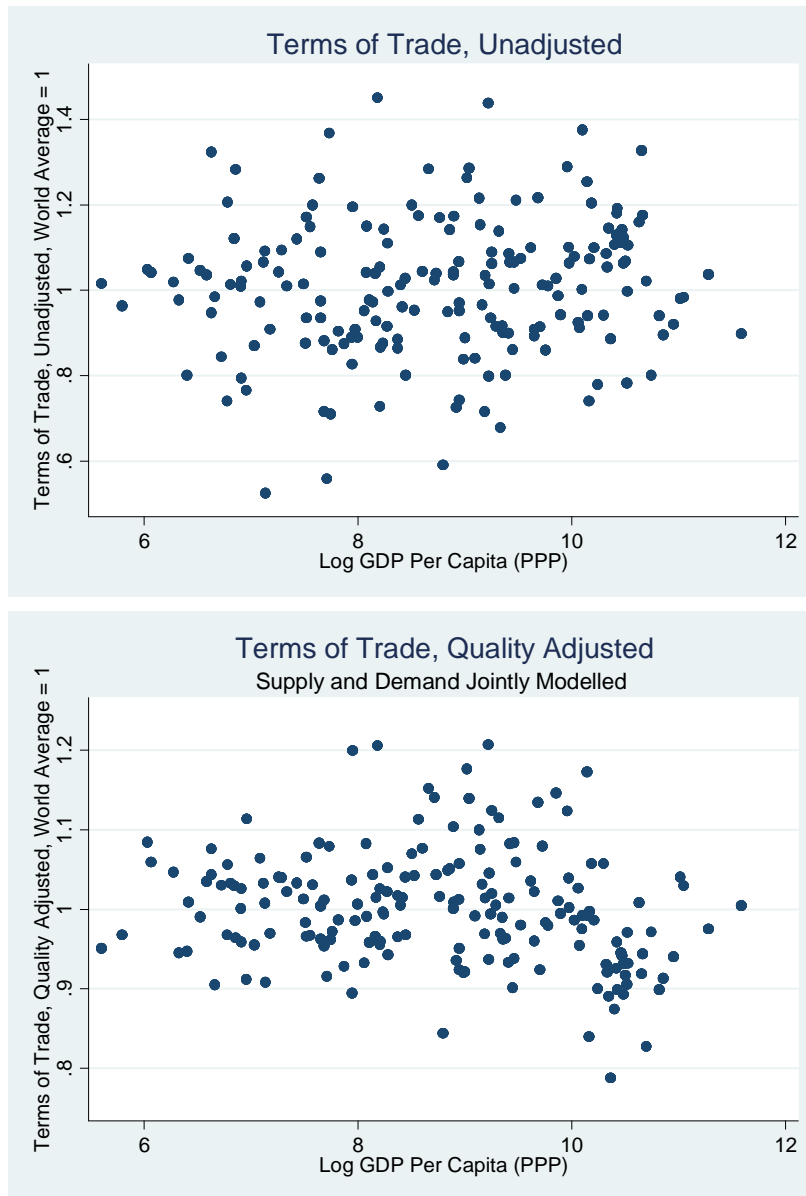
Second, the preference for quality affects import prices in (19), along with bilateral imports X_{igt}^k and total import expenditure Y_{gt}^k , none of which enter the export-side formula in (18). The economic intuition for these terms comes because relative import prices are obtained by comparing a given exporter i selling to two destinations k and l , so that expenditure and tastes of the importer will matter. In our model, any difference in the f.o.b. price from a given exporting firm must be due to quality. As we noted earlier in (10), log quality is only a fraction of the log f.o.b. price, with the remaining difference in f.o.b. prices in (11) attributed to the quality-adjusted price. This pattern is illustrated on the import side in Figure 7.

Terms of Trade

Figure 8 shows terms of trade estimates for 2007. Terms of trade estimates constructed using raw export and import prices fluctuate substantially across countries, and lie between 0.53 and 1.45.²⁸ Terms of trade estimates constructed from quality-adjusted prices move in a much narrower band, between 0.79 and 1.21.²⁹ Notably, the terms of trade decline in real GDP per capita, as wealthier countries are trading higher-quality goods at higher quality-adjusted prices,

²⁸ 0.53 and 1.89 including St Vincent and the Grenadines. See note 25.

²⁹ 0.79 and 1.34 including St Vincent and the Grenadines. See note 25.

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

but this effect is much stronger for imports than for exports. This result is due in part to the lower unit value of imports and exports for poor countries, which have a greater impact on reducing the quality-adjusted import price in (19) than the adjusted export price in (18), because of the smaller exponent on the c.i.f./f.o.b. ratio on the right of (18). But this result also relies on the supply-side intuition from our model: only more efficient exporters can overcome the fixed costs of

selling to countries with small markets, and these firms sell higher quality. Working against this effect is the mechanism of Arkolakis (2010), whereby smaller markets with lower real expenditure (Y^k/p^k) have their fixed costs reduced in (11), and also the reduced demand for quality in low-income countries. In all years quality-adjusted export prices have a modest and usually insignificant relationship with income, while quality-adjusted import prices are usually positively associated with income, and from the mid-1990s significantly so. The terms of trade are consistently significantly negatively related to income from 1993 onwards.³⁰

We report estimates for aggregate export quality for 1987, 1997 and 2007 in Table 2 for the 52 largest traders measured by their average value of exports from 1984 to 2011. Swiss exports have the highest quality, on average 66% higher than the world average in 2007, followed by Israel and Finland with quality 37 percent higher than the average country. Japan, the U.S. and other wealthy European countries usually have 15 to 30 percent higher export quality than average. Of note are the recent quality increases for several Eastern European countries that have joined the EU, especially those proximate to Germany: Czech Republic, Hungary, Poland and Slovakia. Most wealthy industrial countries also exhibit improving relative quality over the 1987-2007 period. Poor large Asian countries have notably lower quality, with Indian and Chinese export quality respectively 13 percent and 34 percent lower than average levels. Vietnam and Indonesia do little better, with quality lagging average levels in 2007 by 12 and 21 percent respectively.

It is interesting that China's relative export quality appears to have declined despite substantial economic progress. This does not imply that its *absolute* export quality has declined, since other countries may have raised quality. China's substantial exports of relatively low-

³⁰ See Figure 13 and the related discussion below.

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

country	Rank				Normalized Quality, World Average = 1			
	1987	1997	2007	Change	1987	1997	2007	Change
Switzerland	1	1	1	0	1.54	1.61	1.66	0.12
Israel	11	3	2	9	1.18	1.36	1.37	0.19
Finland	5	7	3	2	1.22	1.28	1.37	0.15
Ireland	9	6	4	5	1.19	1.31	1.32	0.13
Austria	2	8	5	-3	1.33	1.27	1.32	-0.02
United Kingdom	12	5	6	6	1.18	1.32	1.31	0.13
Sweden	3	2	7	-4	1.24	1.37	1.28	0.04
Japan	13	4	8	5	1.17	1.33	1.27	0.10
France	8	15	9	-1	1.20	1.20	1.26	0.06
USA	4	13	10	-6	1.22	1.20	1.24	0.02
Denmark	10	9	11	-1	1.18	1.26	1.24	0.06
Germany*	6	12	12	-6	1.21	1.21	1.24	0.02
Australia	16	10	13	3	1.14	1.24	1.23	0.10
New Zealand	7	11	14	-7	1.21	1.23	1.21	0.00
Canada	21	16	15	6	1.05	1.18	1.20	0.14
Norway	14	14	16	-2	1.16	1.20	1.20	0.03
Italy	19	17	17	2	1.11	1.18	1.14	0.02
Netherlands	22	21	18	4	1.05	1.11	1.14	0.09
Belgium	17	19	19	-2	1.14	1.14	1.14	0.00
Portugal	20	24	20	0	1.06	1.05	1.13	0.07
Chile	27	20	21	6	0.98	1.12	1.12	0.15
Spain	31	18	22	9	0.95	1.15	1.11	0.15
Nigeria	28	27	23	5	0.97	1.03	1.10	0.13
Algeria	26	30	24	2	0.99	0.98	1.08	0.09
South Africa	34	28	25	9	0.93	1.02	1.05	0.12
Singapore	24	22	26	-2	1.01	1.09	1.04	0.03
Hungary	43	36	27	16	0.84	0.95	1.03	0.20
Mexico	40	41	28	12	0.89	0.90	1.02	0.13
Saudi Arabia	23	33	29	-6	1.04	0.98	1.00	-0.04
Slovakia*	50	46	30	20	0.73	0.84	1.00	0.26
Colombia	25	23	31	-6	1.00	1.09	0.99	0.00
Czech Rep.*	51	42	32	19	0.73	0.89	0.98	0.25
Argentina	39	26	33	6	0.89	1.04	0.97	0.08
UAE	18	25	34	-16	1.12	1.05	0.96	-0.15
Russia*	45	39	35	10	0.82	0.90	0.96	0.14
Turkey	29	31	36	-7	0.96	0.98	0.95	-0.01
Philippines	42	37	37	5	0.88	0.92	0.95	0.06
Iran	15	32	38	-23	1.14	0.98	0.94	-0.20
Brazil	32	29	39	-7	0.95	1.02	0.94	-0.01
Rep. of Korea	38	34	40	-2	0.89	0.97	0.93	0.04
Romania	47	50	41	6	0.81	0.80	0.92	0.11
Malaysia	36	47	42	-6	0.91	0.84	0.90	-0.01
Poland	52	44	43	9	0.71	0.87	0.89	0.18
Thailand	44	45	44	0	0.83	0.86	0.89	0.06
Venezuela	35	43	45	-10	0.93	0.88	0.89	-0.04
Viet Nam	41	35	46	-5	0.89	0.96	0.88	-0.01
India	33	38	47	-14	0.93	0.92	0.87	-0.06
Taiwan	48	48	48	0	0.80	0.83	0.87	0.06
Hong Kong	37	49	49	-12	0.90	0.81	0.86	-0.04
Ukraine*	46	51	50	-4	0.82	0.75	0.79	-0.02
Indonesia	30	40	51	-21	0.95	0.90	0.79	-0.17
China	49	52	52	-3	0.78	0.69	0.66	-0.12
Mean:					1.01	1.06	1.07	
Standard Deviation:					0.17	0.19	0.19	

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

quality products may have in fact caused most other countries to focus on higher quality goods; see Amiti and Khandelwal (2009) for a discussion. We can find plenty of examples in the detailed data of rising relative quality for China, such as “Computers” rising from 0.37 in 1987 to 0.45 in 1997 and 0.75 in 2007, or “Coarse Ceramic Housewares” (dinnerware), rising from 0.40 in 1987 and 1997 to 0.49 in 2007, or “Footwear”, rising from 0.30 in 1987 to 0.57 in 1997 and 0.87 in 2007. But there are an almost equal number of examples of falling relative quality. At the SITC 4-digit level the median quality estimate for China has risen modestly from 0.58 in 1987 to 0.59 in 1997 and 0.62 in 2007. What is working against China in aggregate are the weights applied to items due to compositional shifts in China’s exports. In 1987, 62 percent of China’s exports were in BEC categories 1 through 3: Food, Industrial Supplies, and Fuels. China’s measured quality was much closer to average levels for these products, varying from 0.87 for Industrial Supplies to 0.94 for Fuels. By 1997 these exports had declined to 35 percent of China’s exports, and to just 27 percent by 2007. China’s exports at first were mostly re-oriented towards consumer goods (BEC 6), with that share rising from 30 percent in 1987 to 44 percent in 1997, but these declined back to 27 percent in 2007. The more prolonged re-orientation was towards capital goods and parts (BEC 4), rising from 3 percent of China’s exports in 1987 to 17 percent in 1997 and 39 percent in 2007. It is in capital goods and parts where China’s relative export quality has always been lowest, between 38 and 52 percent of average levels. China’s re-allocation from sectors of relatively high quality towards sectors with relatively low quality is also helping to mask the quality improvements that we often observe as consumers.

Tables 3 through 8 report export quality results for the top-20 exporters in each 1-digit Broad Economic Category (BEC). With a few notable exceptions, the pattern for aggregate quality holds in each of the BEC categories: rich countries tend to have high quality in all BEC

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

country	Rank				Normalized Quality, World Average = 1			
	1987	1997	2007	Change	1987	1997	2007	Change
Australia	8	2	1	7	1.14	1.26	1.41	0.26
Ireland	5	4	2	3	1.18	1.22	1.38	0.20
France	3	3	3	0	1.22	1.24	1.38	0.17
United Kingdom	1	1	4	-3	1.25	1.27	1.37	0.12
Denmark	4	5	5	-1	1.20	1.19	1.20	0.00
New Zealand	10	9	6	4	1.11	1.13	1.19	0.09
Italy	9	10	7	2	1.14	1.10	1.18	0.04
USA	6	6	8	-2	1.16	1.17	1.18	0.02
Germany*	7	11	9	-2	1.16	1.09	1.17	0.01
Netherlands	2	14	10	-8	1.22	1.05	1.16	-0.06
Belgium	12	7	11	1	1.02	1.15	1.15	0.13
Canada	11	8	12	-1	1.06	1.14	1.13	0.07
Spain	16	15	13	3	0.93	1.04	1.08	0.15
Malaysia	17	12	14	3	0.90	1.07	1.00	0.10
Mexico	15	19	15	0	0.93	0.88	0.99	0.07
Brazil	14	17	16	-2	0.99	1.02	0.93	-0.06
Thailand	20	16	17	3	0.76	1.02	0.89	0.13
Argentina	19	20	18	1	0.81	0.87	0.86	0.05
Indonesia	13	13	19	-6	1.00	1.06	0.84	-0.15
China	18	18	20	-2	0.88	0.97	0.82	-0.06
Mean:					1.05	1.10	1.12	
Standard Deviation:					0.15	0.11	0.19	

(* 1987 data are from West Germany)

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

country	Rank				Normalized Quality, World Average = 1			
	1987	1997	2007	Change	1987	1997	2007	Change
Switzerland	1	2	1	0	1.66	1.56	1.56	-0.09
Japan	2	1	2	0	1.56	1.67	1.54	-0.02
United Kingdom	5	3	3	2	1.29	1.40	1.35	0.06
Sweden	4	4	4	0	1.34	1.30	1.30	-0.04
Austria	3	7	5	-2	1.37	1.23	1.30	-0.08
USA	9	6	6	3	1.23	1.24	1.29	0.06
France	8	9	7	1	1.24	1.22	1.27	0.04
Germany*	6	10	8	-2	1.26	1.18	1.26	0.01
Hong Kong	11	5	9	2	1.12	1.26	1.24	0.12
Italy	7	8	10	-3	1.24	1.23	1.22	-0.02
Netherlands	12	16	11	1	1.12	1.07	1.20	0.08
Canada	16	12	12	4	1.03	1.16	1.17	0.14
Spain	17	15	13	4	1.02	1.08	1.15	0.12
Australia	10	13	14	-4	1.17	1.11	1.14	-0.03
Rep. of Korea	13	11	15	-2	1.11	1.17	1.13	0.01
Belgium	14	17	16	-2	1.09	1.05	1.11	0.01
Taiwan	15	14	17	-2	1.09	1.11	1.08	-0.01
Brazil	18	18	18	0	0.98	1.00	0.98	0.01
Russia*	20	19	19	1	0.75	0.87	0.96	0.22
China	19	20	20	-1	0.87	0.84	0.78	-0.09
Mean:					1.18	1.19	1.20	
Standard Deviation:					0.21	0.20	0.18	

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

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

country	Rank				Normalized Quality, World Average = 1			
	1987	1997	2007	Change	1987	1997	2007	Change
United Kingdom	4	3	1	3	1.13	1.19	1.17	0.04
USA	2	2	2	0	1.18	1.20	1.13	-0.05
Netherlands	6	8	3	3	1.02	1.05	1.10	0.08
Saudi Arabia	8	12	4	4	1.00	0.97	1.09	0.10
UAE	9	6	5	4	0.99	1.09	1.04	0.05
Australia	3	1	6	-3	1.18	1.21	1.03	-0.15
Oman	1	10	7	-6	1.21	1.04	1.02	-0.20
Malaysia	5	4	8	-3	1.09	1.11	1.01	-0.08
Nigeria	7	7	9	-2	1.02	1.05	0.99	-0.03
Algeria	14	14	10	4	0.94	0.95	0.96	0.02
Iraq	19	11	11	8	0.87	0.99	0.94	0.07
Russia*	10	16	12	-2	0.98	0.93	0.93	-0.05
Venezuela	11	20	13	-2	0.96	0.85	0.92	-0.03
Norway	12	5	14	-2	0.95	1.09	0.92	-0.03
Canada	15	9	15	0	0.93	1.04	0.90	-0.04
Indonesia	17	13	16	1	0.91	0.96	0.87	-0.04
Mexico	13	15	17	-4	0.95	0.93	0.87	-0.07
Qatar	20	17	18	2	0.86	0.89	0.87	0.01
Iran	16	19	19	-3	0.93	0.86	0.86	-0.07
Kuwait	18	18	20	-2	0.90	0.88	0.81	-0.09
Mean:					1.00	1.01	0.97	
Standard Deviation:					0.11	0.11	0.10	

(* 1987 data are from USSR)

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

country	Rank				Normalized Quality, World Average = 1			
	1987	1997	2007	Change	1987	1997	2007	Change
Switzerland	1	1	1	0	1.80	1.93	1.67	-0.13
Canada	2	4	2	0	1.39	1.40	1.34	-0.05
Ireland	6	3	3	3	1.18	1.41	1.30	0.12
Sweden	4	2	4	0	1.22	1.53	1.29	0.07
Germany*	7	7	5	2	1.18	1.25	1.22	0.05
USA	3	5	6	-3	1.24	1.28	1.22	-0.03
United Kingdom	8	6	7	1	1.08	1.28	1.18	0.10
Belgium	10	9	8	2	1.01	1.21	1.12	0.11
France	5	10	9	-4	1.20	1.21	1.11	-0.09
Netherlands	9	11	10	-1	1.07	1.16	1.11	0.03
Japan	11	8	11	0	0.98	1.22	1.10	0.12
Singapore	13	13	12	1	0.89	0.99	1.02	0.13
Italy	16	12	13	3	0.87	1.00	0.99	0.11
Mexico	14	15	14	0	0.88	0.76	0.96	0.08
Malaysia	15	17	15	0	0.88	0.71	0.91	0.03
Rep. of Korea	18	14	16	2	0.57	0.80	0.89	0.32
Thailand	12	16	17	-5	0.92	0.74	0.82	-0.10
Taiwan	20	18	18	2	0.49	0.57	0.74	0.25
Hong Kong	17	19	19	-2	0.61	0.56	0.71	0.10
China	19	20	20	-1	0.52	0.38	0.52	0.00
Mean:					1.00	1.07	1.06	
Standard Deviation:					0.32	0.38	0.26	

(* 1987 data are from West Germany)

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

country	Rank				Normalized Quality, World Average = 1			
	1987	1997	2007	Change	1987	1997	2007	Change
Sweden	3	1	1	2	1.37	1.58	1.38	0.01
United Kingdom	1	3	2	-1	1.56	1.40	1.32	-0.24
Canada	12	4	3	9	0.92	1.32	1.32	0.40
Austria	5	2	4	1	1.28	1.40	1.32	0.04
Spain	9	12	5	4	1.15	1.08	1.30	0.15
USA	2	6	6	-4	1.40	1.27	1.26	-0.14
Belgium	4	10	7	-3	1.30	1.13	1.24	-0.06
Hungary	6	14	8	-2	1.27	0.93	1.23	-0.05
Germany*	8	7	9	-1	1.19	1.25	1.23	0.04
France	18	5	10	8	0.70	1.27	1.20	0.50
Italy	7	8	11	-4	1.20	1.21	1.20	-0.01
Japan	11	11	12	-1	0.97	1.11	1.14	0.17
Mexico	10	16	13	-3	1.10	0.90	1.14	0.04
Netherlands	17	13	14	3	0.83	1.07	1.13	0.30
Czech Rep.*	16	17	15	1	0.84	0.89	1.06	0.21
Brazil	13	9	16	-3	0.92	1.14	1.03	0.11
Poland	20	19	17	3	0.53	0.82	0.91	0.38
Taiwan	15	18	18	-3	0.85	0.83	0.87	0.02
Rep. of Korea	14	15	19	-5	0.86	0.91	0.86	0.00
China	19	20	20	-1	0.67	0.68	0.77	0.10
Mean:					1.05	1.11	1.15	
Standard Deviation:					0.28	0.23	0.18	

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

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

country	Rank				Normalized Quality, World Average = 1			
	1987	1997	2007	Change	1987	1997	2007	Change
Switzerland	1	1	1	0	2.02	2.42	2.63	0.60
Japan	5	2	2	3	1.47	1.87	2.01	0.54
United Kingdom	7	5	3	4	1.38	1.49	1.60	0.21
France	2	3	4	-2	1.70	1.65	1.58	-0.12
Italy	6	4	5	1	1.44	1.53	1.47	0.03
USA	8	9	6	2	1.35	1.33	1.44	0.09
Ireland	3	6	7	-4	1.49	1.48	1.43	-0.06
Belgium	9	8	8	1	1.31	1.39	1.42	0.11
Germany*	4	7	9	-5	1.48	1.46	1.40	-0.08
Canada	12	12	10	2	1.18	1.26	1.37	0.19
Netherlands	11	10	11	0	1.20	1.31	1.34	0.14
Singapore	13	13	12	1	1.05	1.11	1.30	0.25
Spain	10	11	13	-3	1.21	1.26	1.22	0.01
Hong Kong	16	17	14	2	0.85	0.85	1.06	0.20
Rep. of Korea	17	14	15	2	0.85	0.94	1.05	0.20
Mexico	19	16	16	3	0.74	0.92	1.02	0.28
Turkey	14	15	17	-3	0.94	0.93	0.96	0.02
Taiwan	18	18	18	0	0.77	0.83	0.81	0.04
India	15	19	19	-4	0.87	0.80	0.77	-0.10
China	20	20	20	0	0.62	0.66	0.63	0.02
Mean:					1.20	1.27	1.33	
Standard Deviation:					0.36	0.42	0.45	

(* 1987 data are from West Germany)

categories, while poor countries tend to have notably lower quality. The main exceptions are in Table 5 for BEC 3: Fuels and Lubricants, where there is a less clear relationship between export quality and the exporter's level of development. The recent improvement in Eastern European quality is very apparent in their transport equipment exports. China's declining aggregate relative quality also appears in BEC 1: Food and Beverages and BEC 2: Industrial Supplies.

Our export quality estimates call out for a comparison with the quality estimates of Hallak and Schott (2011) and Khandelwal (2010).³¹ We do this in Figure 9 using data from Table IV of Hallak and Schott and in Figure 10 using the median of HS 10-digit quality results for manufactured products generously provided by Amit Khandelwal. We take logs of our Table 2 results to make them more comparable with Hallak-Schott and demean all series.³² Figure 9 compares our normalized quality estimates with Hallak-Schott in 1997 for the forty countries common to all three papers.³³ The correlation is very high, at 0.67, but there is a considerable difference in the dispersion of the two sets of estimates. The standard deviation of the Hallak-Schott quality estimates is 0.45, compared with 0.18 for our matching estimates. The lower dispersion of our estimates partly reflects the “tighter” solution we get for exporter quality by exploiting the supply-side of our model, but may also be due to using world-wide trade data in all products rather than just U.S. manufacturing imports, and different aggregation procedures.

Figure 10 provides the equivalent comparison with Khandelwal (2010). The correlation between the two sets of estimates is lower, at 0.49, and the higher dispersion of Khandelwal's estimates (the standard deviation is 0.77) cannot be directly compared with the other estimates.³⁴ The lower correlation of our estimates with Khandelwal (2010) is primarily driven by different

³¹ Hallak and Schott (2011) and Khandelwal (2010) do not estimate import quality.

³² Khandelwal's quality estimates are not as directly comparable, since if translated to a CES framework they confound quality and the sensitivity of demand to price: see equation 15 of Khandelwal (2010).

³³ Hallak and Schott's quality estimates are linear trends, so it is a simple matter to back out the implied 1997 results.

³⁴ See note 32.

Figure 9: Comparison with Hallak-Schott (2011)

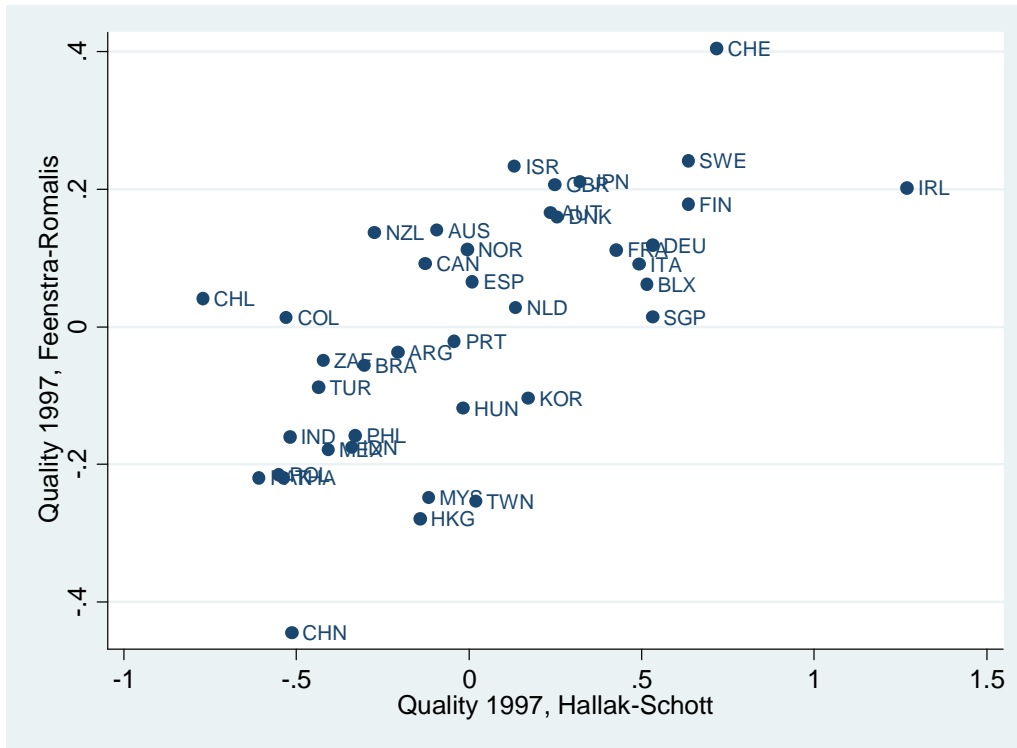


Figure 10: Comparison With Khandelwal (2010)

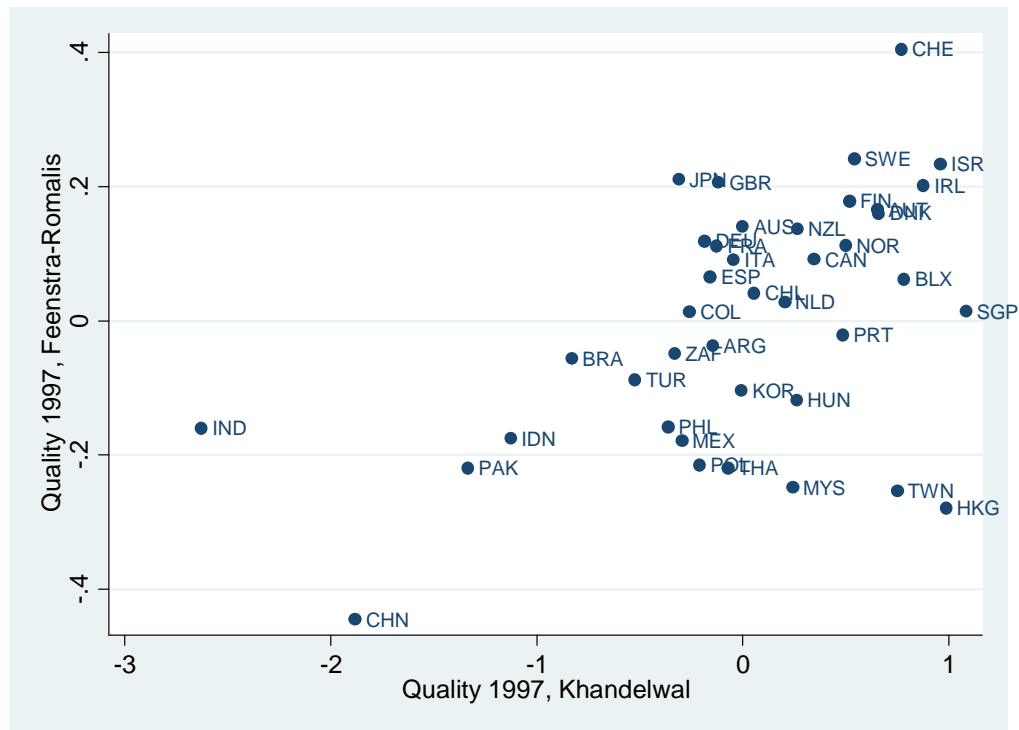


Figure 11: Comparison of Demand-Side Estimates With Khandelwal (2010)

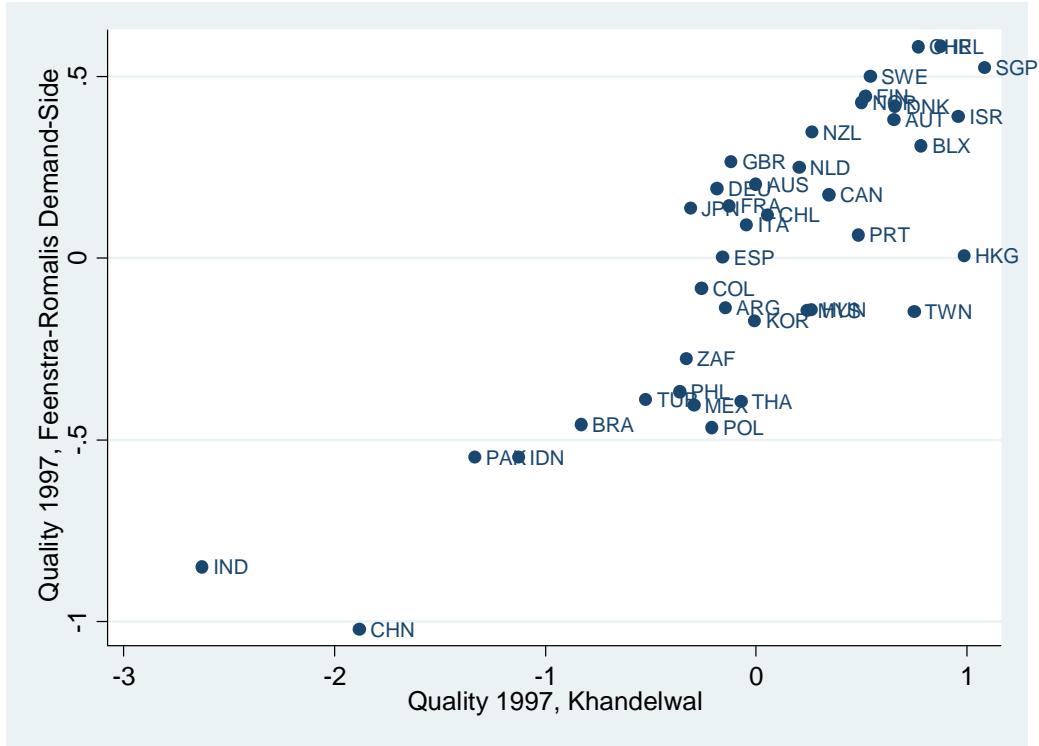
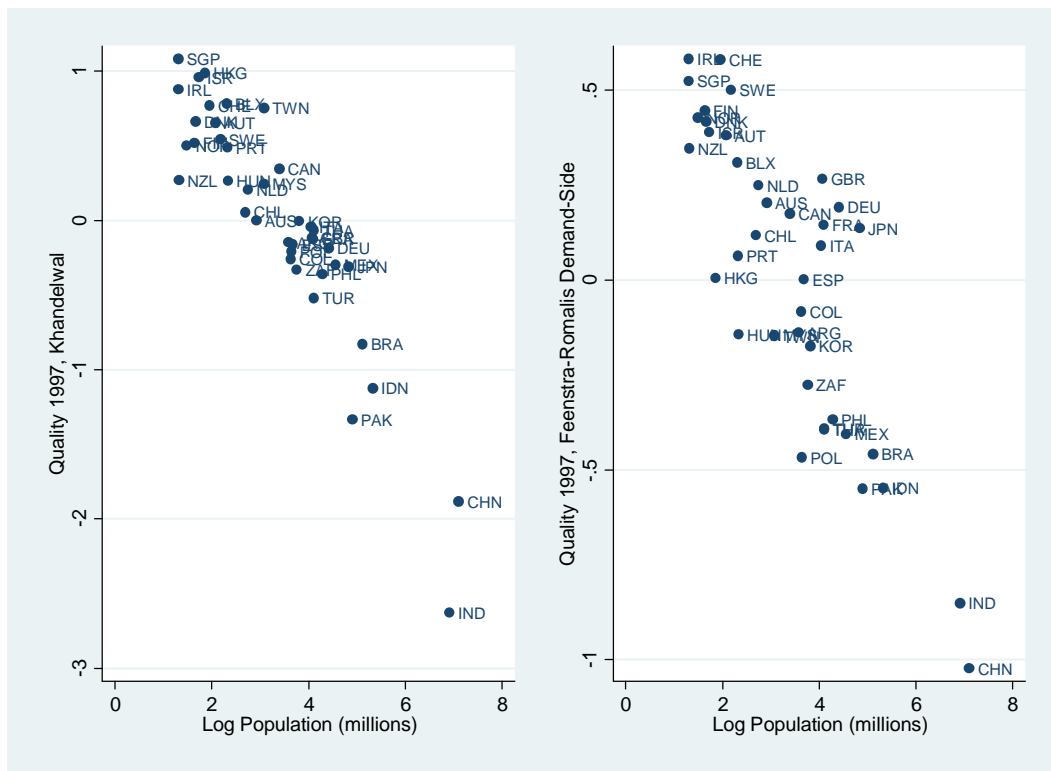


Figure 12: Demand-Side Estimates and Proxy for Number of Firms



supply-side assumptions. We implicitly solve our model for the equilibrium number of firms consistent with observed trade values, while Khandelwal (2010) uses country-population as a proxy of the number of exporting firms.³⁵ In Figure 11 we compare Khandelwal (2010) to our purely illustrative “demand-side” estimates where we also used population as the proxy for the number of exporting firms. The correlation is extremely high at 0.83. Since we use different trade data (world-wide rather than just US imports) and different aggregation methods, the different demand-systems can only be contributing a modest amount to the overall differences in our estimates from Khandelwal (2010).

Figure 12 reveals that these last two sets of estimates – from Khandelwal and our demand-side-only estimates – are extremely negatively correlated with population; the proxy for the number of firms. Less obviously, the Hallak-Schott estimates are closely related to the manufacturing trade balance, which is a key component of their measure of demand. These associations are made crystal-clear in Table 9, which reports regressions of three sets of export quality estimates (Hallak-Schott (2011), Khandelwal (2010) and our “full-model” estimates) plus our import-quality and terms of trade estimates on three country-level variables: log per capita income from PWT; log population; and the manufacturing trade balance from Comtrade divided by manufacturing value added from the World Bank’s World Development Indicators.³⁶

All three export quality estimates are strongly positively correlated with per capita income. Khandelwal’s estimates exhibit a very strong relationship to country population, while Hallak and Schott’s estimates are moderately correlated with population and our estimates

³⁵ Following Khandelwal (2010), we have used the estimated labor force in each SITC industry and country as a proxy for export variety, as explained beneath equation (16). While this proxy enters into the gravity equation (22), and thereby affects the estimated parameters from this equation, it does not otherwise enter into the formulas for quality or quality-adjusted prices.

³⁶ Since Hallak and Schott report trend values of quality, we take an average of the manufacturing trade balance to value added ratio over their 1989 to 2003 sample period.

Table 9: Comparison of Quality Estimates for 1997

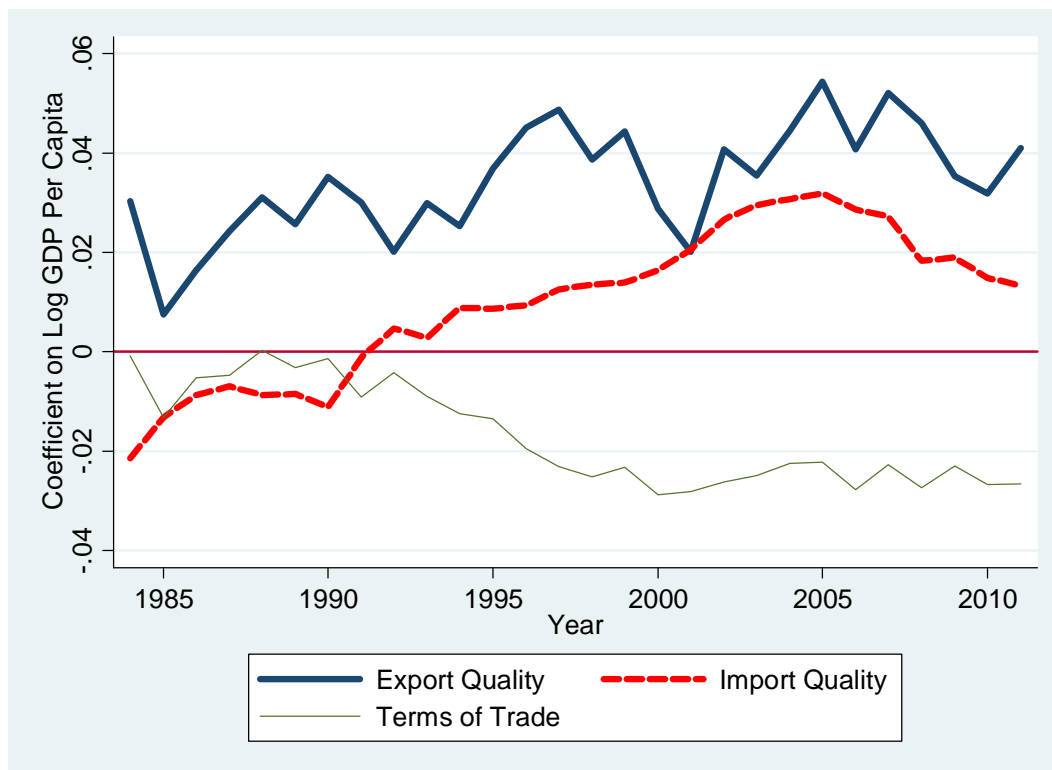
Dependent variable:	Hallak and Schott (2011)	Khandelwal (2010)	Feenstra and Romalis (this paper)		
	<i>Export quality</i>	<i>Export quality</i>	<i>Export quality</i>	<i>Import quality</i>	<i>Terms of trade</i>
Independent variables:					
<i>Log GDP Per Capita</i>	0.32 (0.05)	0.30 (0.07)	0.14 (0.04)	0.02 (0.02)	-0.06 (0.01)
<i>Log Population</i>	-0.08 (0.03)	-0.37 (0.04)	-0.01 (0.02)	-0.01 (0.01)	-0.03 (0.01)
<i>Manufacturing Trade Balance / Value Added</i>	0.84 (0.08)	0.18 (0.11)	0.06 (0.06)	0.02 (0.04)	-0.01 (0.02)
Observations	38	38	38	38	38
R-squared	0.88	0.92	0.52	0.20	0.43

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

(derived from using both the demand and supply side) are uncorrelated with population. The Hallak-Schott quality estimates are very strongly correlated with the manufacturing trade balance, while Khandelwal's and our export quality estimates are only slightly correlated with that balance. Our import quality estimates are not significantly correlated with any of the three variables. Finally, our quality-adjusted terms of trade estimates for these countries are negatively correlated with per capita income and population, but are not associated with the manufacturing trade balance. The key lesson we take from these comparisons is that estimates for quality are very sensitive to proxies chosen for important model variables, whether it be population as the proxy for the number of firms or the manufacturing trade balance as a measure of demand. We have reduced our sensitivity to such proxies by more fully exploiting the supply-side structure of our heterogeneous-firms model, to simultaneously solve for the quality-adjusted prices and (implicitly) the number of firms that are consistent with observed trade data.

We repeat the Table 9 regressions on our export quality, import quality and terms of trade results for each year, using the full sample of countries. Each coefficient on log GDP per capita is plotted in Figure 13. Both export quality and import quality have become more positively associated with income over time, though the prolonged recession in much of the developed world may be eroding the relationship for imports from 2008. The coefficient for exports almost always lies above that for imports, suggesting that richer countries tend to be net exporters of higher quality products, consistent with the proposition of Fajgelbaum, Grossman and Helpman (2011a). Their model generates this result because the production of high quality goods occurs in high-income countries, where demand is greatest. We have a different supply-side mechanism at work, whereby only the most efficient exporters can cover the fixed costs of selling to countries

Figure 13: Coefficients on Log GDP Per Capita



with low import volumes (because they are poor or simply small), and these efficient exporters sell higher quality. The terms of trade become significantly negatively associated with income from 1993.

6. Conclusions

Our goal has been to adjust observed trade unit values for quality so as to estimate quality-adjusted prices in trade. We achieve this goal by explicitly modeling the quality choice by exporting firms in an environment where consumers have non-homothetic tastes for quality. We find a greater preference for quality in richer countries, consistent with Hallak (2006). Our key parameter estimate of the elasticity of quality with respect to the quantity of inputs almost always lies between zero and unity, as required by our model. This implies that only a fraction of observed import unit-value differences are due to quality, with the remainder reflecting differences in quality-adjusted import prices. A key advantage we gain from more fully exploiting the supply-side structure of a heterogeneous firms model is that we reduce our reliance on proxies for some critical features of our model, notably the number of firms. Instead of arbitrarily choosing a proxy, we implicitly solve for the number of firms consistent with our model and observed trade values.³⁷

Our estimates of the elasticity of substitution between different varieties of the same SITC 4-digit products are substantially higher than in Broda and Weinstein (2006). As a result, the observed differences in export unit-values are attributed predominantly to quality, with very small remaining differences in quality-adjusted export prices. The quality-adjusted terms of trade therefore declines with country income in all years since 1993, reflecting rich countries'

³⁷ We have not eliminated our reliance on such proxies, which do indirectly affect quality estimates through their impact on parameter estimates and through our fixed export cost estimates. See note 35.

preferences for higher quality and therefore higher quality-adjusted prices. In that year variation in the quality-adjusted terms of trade is only one-half as large as that in the unadjusted ratio of export to import unit-value indexes.

There are at least two directions for further research. First, as we have noted, our results lend support to the proposition of Fajgelbaum, Grossman and Helpman (2011a) that poor countries are net importers of high-quality goods. They argue that such a trade pattern will disproportionately benefit wealthy consumers in poor countries. It would likewise be of interest to empirically examine this. Our detailed SITC 4-digit estimates of import prices and quality could be used to compute the impact of trade openness on consumers of different income groups, thereby showing how trade interacts with the income distribution of countries.

Second, our finding that the quality-adjusted terms of trade are declining with the level of development give only a partial view on country welfare, and should be combined with the impact of import variety on welfare. Hummels and Klenow (2005) argue that import variety is greater for wealthier countries, and Feenstra (2010) shows how this effect leads to a *positive* relationship between variety-adjusted terms of trade and GDP per capita. Both the quality and the variety effects should be combined to obtain a more complete view of the impact of trade on countries at different levels of income.

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Online Appendix to International Prices and Endogenous Quality

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August 2013

Appendix A: Alternative Cost Specification

Kugler and Verhoogen (2012, p. 321) consider a production function for quality of the form $z = (\varphi^\theta + \psi^\theta)^{1/\theta}$, where φ denotes the quality of inputs used and ψ denotes the plant capability. To relate this production function to our model, we introduce the notation for source country i , firm j and destination country k , and we begin by replacing φ_{ij} by $\varphi_{ij} l_{ij}^k$, indicating the effective amount of the composite input per unit of output needed to produce quality z_{ij}^k . The CES structure of the quality production function in Kugler and Verhoogen does not allow us to obtain a log-linear solution for quality. Instead, we adopt the simpler functional form:

$$z_{ij}^k = (\varphi_{ij} l_{ij}^k + \psi_{ij}^k)^\theta . \quad (\text{A1})$$

If $\psi_{ij}^k > 0$ in (A1) then we continue to interpret this magnitude as the plant capability in producing quality for market k . But we also allow for $\psi_{ij}^k < 0$, in which case we interpret this magnitude as the amount of the composite input needed to produce each unit – regardless of quality – so that $\varphi_{ij} l_{ij}^k + \psi_{ij}^k > 0$ is the “net” amount of the composite input devoted to upgrading quality.

With the quality production function in (A1), the cost function for the firm is:

$$c_{ij}(z_{ij}^k, w_i) = w_i l_{ij}^k = w_i [(z_{ij}^k)^{1/\theta} - \psi_{ij}^k] / \varphi_{ij} . \quad (\text{A2})$$

We can solve for the optimal quality choice from the first-order condition (5), obtaining:

$$\ln z_{ij}^k = \theta \left\{ \ln [T_i^k (\varphi_{ij} / w_i) - \psi_{ij}^k] + \ln [\alpha^k \theta / (1 - \alpha^k \theta)] \right\}, \quad (\text{A3})$$

where for an interior solution we require $T_i^k (\varphi_{ij} / w_i) > \psi_{ij}^k$, meaning that the specific transport costs are large enough as compared to plant capability. The first-order condition for the f.o.b. price is unchanged:

$$(p_{ij}^{*k} + T_i^k) = [c_{ij}(z_{ij}^k, w_i) + T_i^k] \left(\frac{\sigma}{\sigma - 1} \right). \quad (\text{A4})$$

Combining (A2)–(A4), we can solve for the f.o.b. price inclusive of specific transport costs:

$$(p_{ij}^{*k} + T_i^k) = \left(\frac{\sigma}{\sigma - 1} \right) \left(\frac{1}{1 - \alpha^k \theta} \right) \left[T_i^k - \psi_{ij}^k \left(\frac{w_i}{\varphi_{ij}} \right) \right]. \quad (\text{A5})$$

We see from (A4) that the f.o.b. price is increasing in the firm's productivity φ_{ij} if $\psi_{ij}^k > 0$, and is decreasing in productivity if $\psi_{ij}^k < 0$. In the former case, the more productive firm produces substantially higher quality such that the f.o.b. price increases, and in the latter case the quality increase is not that high, so that the price decreases with productivity. With $\psi_{ij}^k = 0$ as we assume in the paper, we have the borderline case where the price is independent of productivity.

Substituting (A5) into (A3), we can solve for quality as:

$$\ln z_{ij}^k = \theta \left\{ \ln [\alpha^k \theta (\sigma - 1) (p_{ij}^{*k} + T_i^k) / \sigma] - \ln (w_i / \varphi_{ij}) \right\}. \quad (\text{A6})$$

Comparing (A6) with (8), we see that a log-linear solution for quality is still obtained and that it is independent of the unobserved variable ψ_{ij}^k . But instead of the f.o.b. price appearing on the right as in (8), it is now the f.o.b. price inclusive of specific transport costs. As mentioned in the text, this variable lies strictly in-between the f.o.b. and c.i.f. prices (since the latter also includes *ad valorem* transport costs), so it is not observed in the data.

Appendix B: Heterogeneous Firms

We first show how the quality-adjusted prices in our heterogeneous-firm model depend on firm productivity. Using quality in (8), the quality-adjusted price $P_{ij}^k = \overline{p_i^k} / z_{ij}^{\alpha^k}$ is:

$$P_{ij}^k = \overline{p_i^k} \left[(w_i / \varphi_{ij}) / \kappa_1^k \overline{p_i^{*k}} \right]^{\alpha^k \theta}. \quad (\text{B1})$$

Since from (7a) and (7b) 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 productivity φ_{ij} of the exporter. We will assume a Pareto distribution for productivities with parameter γ .

We now solve the heterogenous firm model to obtain the average quality-adjusted prices. To exploit the ZCP condition we integrate over all firms with productivity above the marginal exporter, obtaining total sectoral exports from country i to k . To aggregate over exporters, note that the ratio of demand for firm j and the cutoff firm, 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 from country i to k are:

$$\begin{aligned} X_i^k &\equiv M_i \int_{\hat{\varphi}_i^k}^{\infty} X_{ij}^k g_i(\varphi) d\varphi = M_i \int_{\hat{\varphi}_i^k}^{\infty} \hat{X}_i^k (P_{ij}^k / \hat{P}_i^k)^{1-\sigma} g_i(\varphi) d\varphi \\ &= M_i \hat{X}_i^k \left[\int_{\hat{\varphi}_i^k}^{\infty} \left(\frac{\hat{\varphi}_i^k}{\varphi} \right)^{\alpha^k \theta (1-\sigma)} g_i(\varphi) d\varphi \right] = \hat{X}_i^k M_i \left(\frac{\hat{\varphi}_i^k}{\varphi_i} \right)^{-\gamma} \kappa_2^k, \end{aligned} \quad (\text{B2})$$

with $\kappa_2^k \equiv \gamma / [\gamma - \alpha^k \theta (\sigma - 1)]$ and assuming $\gamma > \alpha^k \theta (\sigma - 1)$. Substituting (B2) for \hat{X}_i^k in (12) and using (14) for fixed costs, we solve for the wage relative to the cutoff productivity:

$$\left(\frac{w_i}{\hat{\varphi}_i^k} \right)^{1+\gamma} = \left(\frac{X_i^k / \sigma \kappa_2^k \text{tar}_i^k \tilde{f}_i^k}{M_i (\varphi_i / w_i)^\gamma} \right), \text{ with } \tilde{f}_i^k \equiv \left(\frac{Y^k}{p^k} \right)^{\beta_0} e^{\beta' F_i^k}. \quad (\text{B3})$$

Substituting this solution for productivity-adjusted wages into (11), we readily obtain the quality-adjusted price for the marginal exporter. Following Melitz (2003), we form the CES averages of the quality-adjusted prices in (B1) by integrating over firms in country i exporting to k :

$$\overline{P_i^k} \equiv \left[\int_{\hat{\varphi}_i^k}^{\infty} P_{ij}^k(\varphi)^{(1-\sigma)} \frac{g_i(\varphi)}{[1 - G_i(\hat{\varphi}_i^k)]} d\varphi \right]^{\frac{1}{1-\sigma}} = \left(\kappa_2^k \right)^{\frac{1}{1-\sigma}} \hat{P}_i^k, \quad (\text{B4})$$

obtained by substituting for $P_{ij}^k(\varphi_{ij})$ from (B1) and computing the integral for $\gamma > \alpha^k \theta (\sigma - 1)$.

This expression shows that the *average* quality-adjusted price $\overline{P_i^k}$ is proportional to the *cut-off* price \hat{P}_i^k , with the factor of proportionality depending on model parameters. Combining (9) with (B2)–(B4), we therefore obtain (13).

To develop the gravity equation, we return to the ZCP condition in (10). While the firm-level sales \hat{X}_i^k are not observed in our data, they equal CES demand from the expenditure function in (1a). That is, $\hat{X}_i^k = (\hat{P}_i^k)^{-(\sigma-1)}[Y^k P^{k(\sigma-1)}]$, where P^k is the exact price index corresponding to the CES expenditure function in (1a). We use (B2) and (B4) to re-express this relation using X_i^k and $\overline{P_i^k}$, and then rearrange terms and substitute for $w_i / \hat{\varphi}_i^k$ using (B3). This gives us (17) in the text. While this completes the derivation of formulas used in the main text, we continue here to derive a gravity equation closer to Chaney (2008).

Substitute (17) into (13) to obtain another expression for the quality-adjusted price $\overline{P_i^k}$:

$$\overline{P_i^k} = \left(\left(\kappa_2^k \right)^{1/(1-\sigma)} \overline{P_i^k} / \left(\kappa_1^k \overline{P_i^{*k}} \right)^{\alpha^k \theta} \right)^{\frac{1}{1+\alpha^k \theta(\sigma-1)}} \left(\frac{Y^k P^{k(\sigma-1)}}{\sigma \kappa_2^k \text{tar}_i^k \tilde{f}_i^k} \right)^{\frac{\alpha^k \theta}{1+\alpha^k \theta(\sigma-1)}}. \quad (\text{B5})$$

To re-express (17) in a form closer to the gravity equation, we need to solve for the CES price index P^k , which is:

$$P^k \equiv \left[\sum_i M_i \int_{\hat{\varphi}_i^k}^{\infty} P_{ij}^k(\varphi)^{(1-\sigma)} g_i(\varphi) d\varphi \right]^{1/(1-\sigma)} = \left[\sum_i \overline{P_i^k}^{(1-\sigma)} M_i \left(\frac{\hat{\varphi}_i^k}{\varphi_i} \right)^{-\gamma} \right]^{1/(1-\sigma)}, \quad (\text{B6})$$

using (B4). We obtain the gravity equation by solving for this CES price index. To this end, we first substitute (B5) into (17) and simplify to obtain exports,

$$\begin{aligned} \frac{X_i^k}{M_i(\varphi_i / w_i)^\gamma} &= \left(\left(\kappa_2^k \right)^{1/(1-\sigma)} \overline{P_i^k} / \left(\kappa_1^k \overline{P_i^{*k}} \right)^{\alpha^k \theta} \right)^{\frac{-(\sigma-1)(1+\gamma)}{1+\alpha^k \theta(\sigma-1)}} \left(\frac{Y^k P^{k(\sigma-1)}}{\sigma \kappa_2^k \text{tar}_i^k \tilde{f}_i^k} \right)^{\frac{-\alpha^k \theta(\sigma-1)(1+\gamma)}{1+\alpha^k \theta(\sigma-1)}} \\ &\quad \times \left(Y^k P^{k(\sigma-1)} \right)^{(1+\gamma)} \left(\sigma \kappa_2^k \text{tar}_i^k \tilde{f}_i^k \right)^{-\gamma}. \end{aligned}$$

Dividing by $\sigma \kappa_2^k \text{tar}_i^k \tilde{f}_i^k$ on both sides and simplifying, we obtain:

$$\left(\frac{X_i^k / \sigma \kappa_2^k \text{tar}_i^k \tilde{f}_i^k}{M_i (\varphi_i / w_i)^\gamma} \right) = \left(\frac{\overline{p}_i^k / \left(\kappa_1^k \overline{p}_i^{*k} \right)^{\alpha^k \theta}}{\left[1 + \alpha^k \theta (\sigma - 1) \right]^{\frac{-(\sigma - 1)(1 + \gamma)}{1 + \alpha^k \theta (\sigma - 1)}}} \left(\frac{Y^k P^{k(\sigma - 1)}}{\sigma \text{tar}_i^k \tilde{f}_i^k} \right)^{\frac{(1 + \gamma)}{1 + \alpha^k \theta (\sigma - 1)}} \right). \quad (\text{B7})$$

Second, we solve for the export probabilities $(\hat{\varphi}_i^k / \varphi_i)^{-\gamma}$ appearing in (B6) using (B3),

$$\left(\frac{\hat{\varphi}_i^k}{\varphi_i} \right)^{-\gamma} = \left(\frac{X_i^k / \sigma \kappa_2^k \text{tar}_i^k \tilde{f}_i^k}{M_i (\varphi_i / w_i)^\gamma} \right)^{\frac{\gamma}{(1 + \gamma)}} \left(\frac{w_i}{\varphi_i} \right)^{-\gamma}. \quad (\text{B8})$$

We now follow the same steps as in Chaney (2008), which means that we substitute (B7) into (B8) 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 (B6) to solve for the CES price index in terms of those other variables. That solution is:

$$Y^k P^{k(\sigma - 1)} = \left(\frac{Y^k}{\kappa_2^k R^k} \right)^{\frac{[1 + \alpha^k \theta (\sigma - 1)]}{(1 + \gamma)}}, \quad (\text{B9})$$

where R^k is a ‘‘remoteness’’ variable defined by,³⁸

$$R^k \equiv \sum_i M_i \left(\frac{\varphi_i}{w_i} \right)^\gamma \left(\frac{\overline{p}_i^k / \left(\kappa_1^k \overline{p}_i^{*k} \right)^{\alpha^k \theta}}{\left[1 + \alpha^k \theta (\sigma - 1) \right]^{\frac{-(\sigma - 1)(1 + \gamma)}{1 + \alpha^k \theta (\sigma - 1)}}} \left(\sigma \text{tar}_i^k \tilde{f}_i^k \right)^{\frac{[\gamma - \alpha^k \theta (\sigma - 1)]}{1 + \alpha^k \theta (\sigma - 1)}} \right). \quad (\text{B10})$$

The gravity equation is obtained by substituting (B9) back into (B7):

$$\left(\frac{X_i^k}{M_i (\varphi_i / w_i)^\gamma} \right) = \left(\frac{\overline{p}_i^k / \left(\kappa_1^k \overline{p}_i^{*k} \right)^{\alpha^k \theta}}{\left[1 + \alpha^k \theta (\sigma - 1) \right]^{\frac{-(\sigma - 1)(1 + \gamma)}{1 + \alpha^k \theta (\sigma - 1)}}} \left(\sigma \text{tar}_i^k \tilde{f}_i^k \right)^{\frac{[\gamma - \alpha^k \theta (\sigma - 1)]}{1 + \alpha^k \theta (\sigma - 1)}} \left(\frac{Y^k}{R^k} \right) \right). \quad (\text{B11})$$

The exponents in this gravity equation appear complex, but in fact, are not too different from those in Chaney (2008) as can be seen by allowing $\alpha^k \theta \rightarrow 1$. In this limit $\kappa_1^k \rightarrow (\sigma - 1) / \sigma$

³⁸ This term is the inverse of the ‘‘remoteness’’ variable derived by Chaney (2008).

which can be ignored as a constant. Then the price term in (B11) approaches the ratio $\overline{(p_{it}^k / p_{it}^{*k})}$, which from (7a) and (7b) equals the iceberg trade costs τ_i^k . The exponent of that term approaches $-(\sigma - 1)(1 + \gamma) / \sigma$. In contrast, Chaney (2008) finds that the exponent of iceberg trade costs is simply the Pareto parameter $-\gamma$. This difference between our gravity equation and Chaney's is explained by the fact that we have allowed the fixed costs of exporting to depend on the productivity of the firm. The second terms on the right of (B11) are the components of the fixed costs of exporting, inclusive of one plus the *ad valorem* tariff, raised to a power that is again similar to that in Chaney (2008) when $\alpha^k \theta \rightarrow 1$.

Lastly, we note that the rate at which the quality-adjusted price in (B1) declines with productivity, or sales expand, differs from that rate in Chaney (2008). For the CES expenditure function in (1a), sales depend on the quality-adjusted price with elasticity $(1 - \sigma)$, and from (B1), the price depends on productivity with elasticity $-\alpha^k \theta$, so that firms' sales depends on productivity with elasticity $\alpha^k \theta (\sigma - 1)$. It follows that firms' sales in our endogenous-quality model are Pareto distributed with parameter $\zeta^k = \gamma / [\alpha^k \theta (\sigma - 1)]$ in country k . In contrast, $\zeta^k = \gamma / (\sigma - 1)$ in Chaney (2008), so our model has the same properties as $\alpha^k \theta \rightarrow 1$.

Appendix C: Data and Calibration

(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 and IDB databases 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;³⁹ (iii) U.S. tariff schedules from the U.S. International Trade Commission from 1989 onwards (Feenstra, Romalis and Schott, 2002); (iv) U.S. tariff schedules derived from detailed U.S. tariff revenue and trade data from 1974 to 1988 maintained by the Center for International Data at UC Davis; and (v) the texts of preferential trade agreements primarily sourced from the WTO's website, the World Bank's Global Preferential Trade Agreements Database, or the Tuck Center for International Business Trade Agreements Database. For the U.S., specific tariffs have been converted into ad-valorem tariffs by dividing by the average unit value of matching imported products. Due to the difficulties of extracting specific tariff information for other countries and matching it to appropriate unit values, only the ad-valorem component of their tariffs are used. The vast 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 program⁴⁰ to the 4-digit SITC Revision 2 level by taking the simple average of tariff lines within each SITC code.

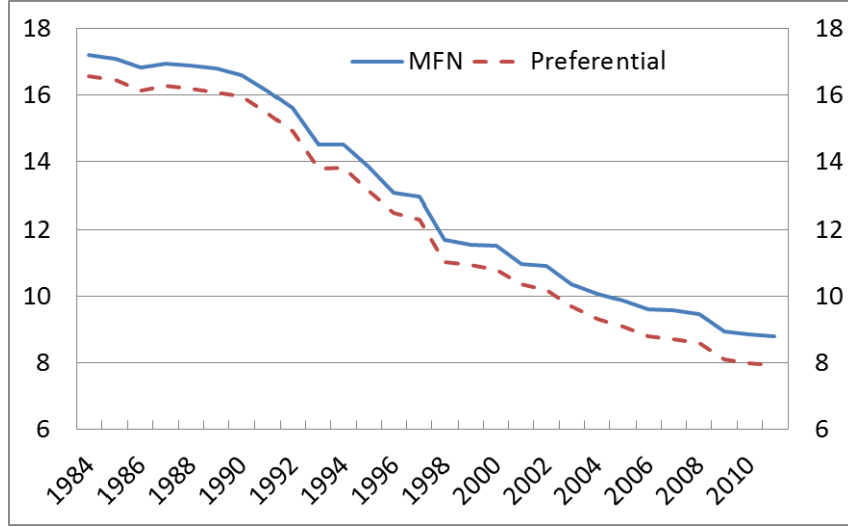
Tariff schedules are often not available in each year, especially for smaller countries. Updated schedules are more likely to be available after significant tariff changes. Rather than replacing “missing” MFN tariffs by linearly interpolating observations, missing observations are set equal to the nearest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

³⁹ Most tariff schedules can be fairly readily matched to the SITC classification.

⁴⁰ 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.

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; and (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 keep the most favorable potentially applicable preferential tariff. Punitive non-MFN tariff levels tend not to change over time (though the countries they apply to do change). We replace missing observations in the same way we replace missing MFN tariff observations. The evolution of a simple average of these MFN and most favorable preferential tariffs from 1984-2011 is shown 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.

(iv) *Quality-Adjusted Unit Values:* The quality estimates shown by (18) and (19) depend on the c.i.f. and f.o.b. unit values, but two-thirds of the bilateral, 4-digit SITC trade flows in our 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. or f.o.b. data. To achieve this we use the structure of our model, where from (7a) and (7b) the ratio of c.i.f. to f.o.b. prices is proportional to the *ad valorem* trade costs, $\overline{p_{igt}^k} / \overline{p_{igt}^{*k}} = \kappa_{3g}^k \tau_{igt}^k$, where

Appendix Figure 1: Typical MFN and Preference-adjusted Tariff 1984-2011*

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

$\kappa_{3g}^k \equiv \left[\left(\frac{1}{1-\alpha_g^k \theta_g} \right) \left(\frac{\sigma_g}{\sigma_g - 1} \right) \right] / \left[\left(\frac{1}{1-\alpha_g^k \theta_g} \right) \left(\frac{\sigma_g}{\sigma_g - 1} \right) - 1 \right]$ and we make explicit the subscripts for goods g and

time t . We use the estimated value of $\hat{\kappa}_{3g}^k$ and the net-of-tariff unit-values from (20), which

means that the *ad valorem* trade costs depend only on distance:

$$\ln(uv_{igt}^k / uv_{igt}^{*k}) = \eta_{gt} + \ln \hat{\kappa}_{3g}^k + \eta_{1g} \ln dist_i^k + u_{igt}^k - u_{igt}^{*k}. \quad (C1)$$

We estimate the parameters η_{gt} and η_{1g} using a median regression, which helps controls for unusual values of the measurement errors appearing as the error in (C1). Then when a 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

$uv_{it}^{*k} \hat{\kappa}_3^k \exp(\hat{\eta}_{gt} + \hat{\eta}_{1g} \ln dist_i^k)$, and when only the c.i.f. unit value is available then we impute the f.o.b. unit value by $uv_{it}^k / [\hat{\kappa}_3^k \exp(\hat{\eta}_{gt} + \hat{\eta}_{1g} \ln dist_i^k)]$.

(v) **Language Data:** Data on 6,909 spoken languages in almost all countries is published in M.

Paul Lewis (2009) and available online at www.ethnologue.com (Ethnologue). We collected data

on the number of speakers in each country of languages that are spoken by 0.5 percent or more

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

From this language data we construct the probabilities that randomly chosen people in two countries share the same: (i) language; (ii) language genus; (iii) language family; and (iv) official language. These probabilities are highly correlated, which sometimes complicates the recovery of fixed cost estimates in Appendix E. To overcome this problem we use factor analysis, whereby we approximate the information contained in these four variables by

constructing two orthogonal “principal factors”. These two principal factors are the language variables we use in our regressions.

(vi) Calibration of β_0 : Eaton, Kortum and Kramarz (2011) regressed the number of firms exporting from France, $\ln N_i^k$, on the log of real manufacturing imports from France across various destination countries, obtaining an elasticity of 0.65.⁴¹ A similar regression on French data is reported by Arkolakis (2010). This regression was repeated in Eaton, Kortum and Sotelo (2012) for Brazil, France, Denmark and Uruguay, yielding an elasticity of 0.71 (or 0.62 with country fixed effects). In a Melitz model with identical fixed costs for all firms exporting to country k , those elasticities measure $(1 - \beta_0)$, suggesting an estimate for β_0 of about 0.35.

In our model, the coefficient of 0.65 linking the number of firms to market size implies an estimate for β_0 of less than 0.35, due to our modeling of fixed costs in (11) as depending on the productivity of the cutoff exporter. To see this, start with (B3) where $N_i^k \equiv M_i(\hat{\varphi}_i^k / \varphi_i)^{-\gamma}$ appearing on the right denotes the number (or mass) of exporters. This number is proportional to total exports divided by those of the cutoff exporter, $N_i^k \propto (X_i^k / \hat{X}_i^k)$. Then substitute the cutoff exports \hat{X}_i^k from (10) into (B2), also substitute for $w_i / \hat{\varphi}_i^k$ using (B3), and simplify (ignoring constants, export-country fixed-effects, and tariffs) to obtain:

$$\left(N_i^k\right)^{(1+\gamma)/\gamma} \propto \left(\frac{X_i^k}{f_i^k}\right) = X_i^k \left(\frac{Y^k}{p^k}\right)^{-\beta_0} e^{-\beta' F_i^k},$$

using (11). If the number of exporters has an elasticity of 0.65 with respect to $\ln Y^k$, then since exports have elasticity of unity it follows that $0.65\left(\frac{1+\gamma}{\gamma}\right) = 1 - \beta_0$, so that $\beta_0 = 1 - 0.65\left(\frac{1+\gamma}{\gamma}\right)$, which is the formula we use to calibrate β_0 . Across our industry estimates, γ ranges from about 2 to a very large number, so we see that β_0 ranges from about zero up to 0.35.

⁴¹ This result is not reported in the published paper, and we thank Jonathan Eaton for informing us of it.

Appendix D: Indexes for Price and Quality

To implement (18) and (19), we use the import c.i.f. unit value inclusive of *ad valorem* tariffs, $uv_{igt}^k tar_{igt}^k$, to replace $\overline{p_{igt}^k}$ as in (20), and the f.o.b. unit value uv_{igt}^{*k} to replace $\overline{p_{igt}^{*k}}$.

Defining average quality as the ratio of the tariff-inclusive unit-values to these tariff-inclusive quality adjusted prices, we obtain the following measures of relative quality:

$$\ln \left(\frac{\overline{z_{igt}^k}}{z_{jgt}^k} \right) = \frac{\alpha_g^k \theta_g}{1 + \alpha_g^k \theta_g (\sigma_g - 1)} \left[(\sigma - 1) \ln \left(\frac{uv_{igt}^k}{uv_{jgt}^k} \right) + \ln \left(\frac{uv_{igt}^{*k}}{uv_{jgt}^{*k}} \right) + \beta'_g (F_i^k - F_j^k) + \sigma_g \ln \left(\frac{tar_{igt}^k}{tar_{jgt}^k} \right) \right], \quad (D1)$$

$$\begin{aligned} \ln \left(\frac{\overline{z_{igt}^k}}{z_{igt}^l} \right) &= \frac{\bar{\alpha}_g \theta_g}{(1 + \gamma_g)} \left[(1 + \gamma_g) \ln \left(\frac{\kappa_{1g}^k uv_{igt}^{*k}}{\kappa_{1g}^l uv_{igt}^{*l}} \right) - \ln \left(\frac{X_{igt}^k / tar_{igt}^k}{X_{igt}^l / tar_{igt}^l} \right) + \beta_{0g} \ln \left(\frac{Y_t^k / p_t^k}{Y_t^l / p_t^l} \right) + \beta'_g (F_i^k - F_i^l) \right] \\ &+ \left[\frac{\bar{\alpha}_g \theta_g}{(1 + \gamma_g)} + \frac{1}{(\sigma - 1)} \right] \ln \left(\frac{\kappa_{2g}^k}{\kappa_{2g}^l} \right). \end{aligned} \quad (D2)$$

(D1) defines the relative export quality of country i selling to j and k , while (D2) defines the import quality for countries k and l purchasing from i . The quality-adjusted unit values are then:

$$\left(\frac{UV_{igt}^{*k}}{UV_{jgt}^{*k}} \right) \equiv \left(\frac{uv_{igt}^{*k}}{uv_{jgt}^{*k}} \right) \Big/ \left(\frac{\overline{z_{igt}^k}}{z_{jgt}^k} \right) \quad \text{and} \quad \left(\frac{UV_{igt}^k}{UV_{igt}^l} \right) \equiv \left(\frac{uv_{igt}^k}{uv_{igt}^l} \right) \Big/ \left(\frac{\overline{z_{igt}^k}}{z_{igt}^l} \right). \quad (D3)$$

Notice that the first of these definitions use the net-of-tariff f.o.b. unit values on the right, to measure the relative export unit values for countries i and j selling to k . The second definition uses the net-of-tariff c.i.f. unit values on the right, to measure the relative import unit values for country i selling to destinations k and l . So for consistency with the national accounts definition of import and export price indexes, we are measuring the quality-adjusted import and export unit value indexes in net-of-tariff terms.

To check the Fisher-Shell approach to evaluating quality when preferences differ across

countries, in place of (D2) we instead use the geometric average preference for quality

$\alpha_g^{kl} \equiv (\alpha_g^k \alpha_g^l)^{1/2}$ of the two countries. In that case the terms κ_{1g}^k and κ_{1g}^l are evaluated with the same average preference α_g^{kl} so they are equal, and likewise for κ_{2g}^k and κ_{2g}^l . Then in place of (D2), relative import quality becomes:

$$\ln \left(\frac{\overline{z_{igt}^k}}{z_{igt}^l} \right) = \frac{\alpha_g^{kl} \theta_g}{(1 + \gamma_g)} \left[(1 + \gamma_g) \ln \left(\frac{uv_{igt}^{*k}}{uv_{igt}^{*l}} \right) - \ln \left(\frac{X_{igt}^k / tar_{igt}^k}{X_{igt}^l / tar_{igt}^l} \right) + \beta_{0g} \ln \left(\frac{Y_t^k / p_t^k}{Y_t^l / p_t^l} \right) + \beta'_g (F_i^k - F_i^l) \right]. \quad (D2')$$

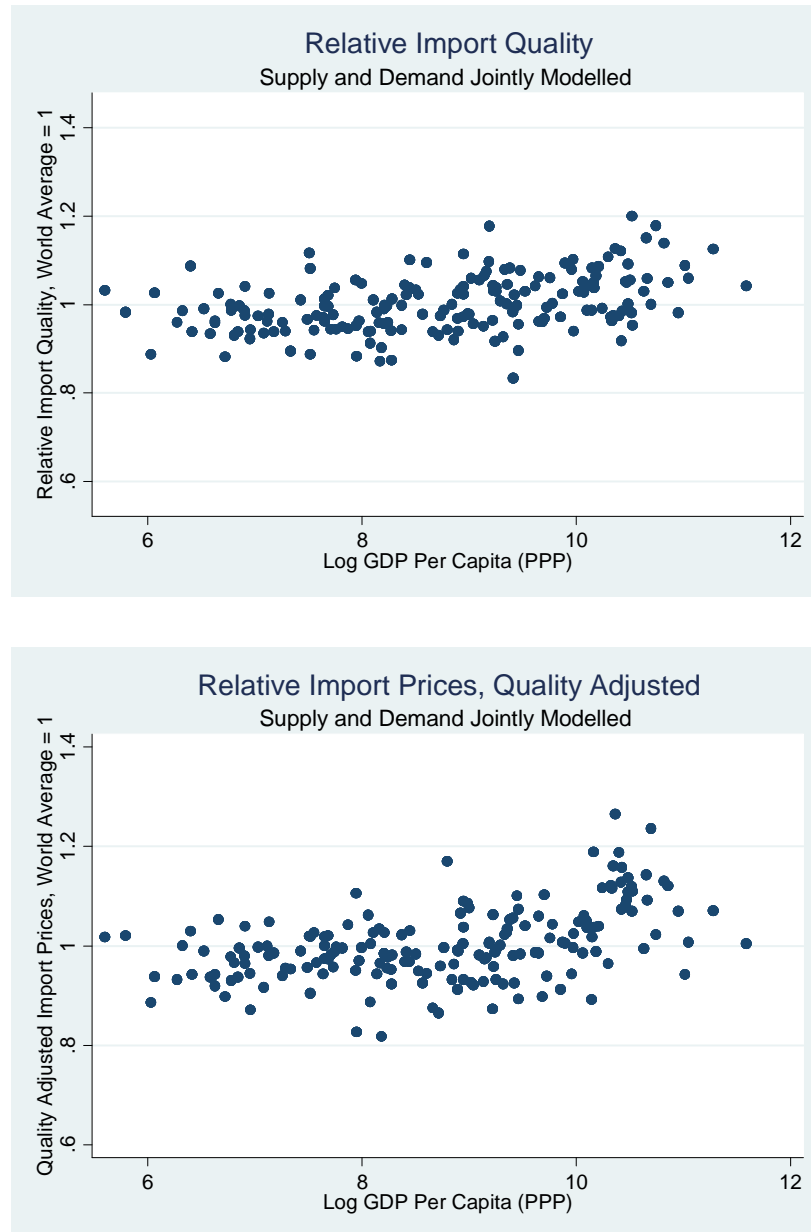
For completeness we also programmed the relative export quality in (D1) using the average preference for quality:

$$\ln \left(\frac{\overline{z_{igt}^k}}{z_{jgt}^k} \right) = \frac{\bar{\alpha}_g \theta_g}{1 + \bar{\alpha}_g \theta_g (\sigma_g - 1)} \left[(\sigma - 1) \ln \left(\frac{uv_{igt}^k}{uv_{jgt}^k} \right) + \ln \left(\frac{uv_{igt}^{*k}}{uv_{jgt}^{*k}} \right) + \beta'_g (F_i^k - F_j^k) + \sigma_g \ln \left(\frac{tar_{igt}^k}{tar_{jgt}^k} \right) \right]. \quad (D1')$$

In the results, we find that methods (D1) and (D1') on relative export quality make only a very minor difference to the results: in both cases we obtain export quality that closely matches the export unit values, leading to quality-adjusted export prices that are quite similar across countries. Method (D1') gives very slightly less variation in quality-adjusted export prices and correspondingly higher variation in export quality, but plots of results are visually indistinguishable from method (D1). Method (D2') on the import side leads to slightly greater variation in quality-adjusted import prices and less variation in estimated import quality than method (D2), but plots of results are very hard to distinguish from Figure 7. In Appendix Figure 2 we report import quality (top panel) and quality-adjusted prices (bottom panel), both panels being almost the same as the corresponding panels in Figure 7.

As noted in the main text, we use a two-stage aggregation procedure that arises naturally from our trade data. In the first stage, for each 4-digit SITC product g we aggregate over all

**Appendix Figure 2: Estimates of Import Quality and Quality Adjusted Import Prices
Using Fisher-Shell Approach, 2007**



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 f.o.b. unit-value ratio ($uv_{git}^{*k} / uv_{gjt}^{*k}$) compares countries i and j selling to k , from we shall construct an

index of *relative export prices*. That is, we compare the unit values of countries i and j only when they are selling to the same country k : essentially, we are treating products sold to different countries as entirely different goods and avoid comparing their prices in that case.

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

$$P_{ijgt}^L \equiv \frac{\sum_{k=1}^{C_{ij}} uv_{igt}^{*k} q_{jgt}^k}{\sum_{k=1}^{C_{ij}} uv_{jgt}^{*k} q_{jgt}^k}, \quad \text{and,} \quad P_{ijgt}^A \equiv \frac{\sum_{k=1}^{C_{ij}} uv_{igt}^{*k} q_{igt}^k}{\sum_{k=1}^{C_{ij}} uv_{jgt}^{*k} q_{igt}^k}. \quad (\text{D4})$$

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

Alternatively, we could instead use the *quality-adjusted* unit values UV_{igt}^{*k} in these formulas, in which case the quantities are instead Q_{igt}^k with $uv_{igt}^{*k} q_{igt}^k = UV_{igt}^{*k} Q_{igt}^k$ 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. Letting $s_{jgt}^{*k} = uv_{jgt}^{*k} q_{jgt}^k / \sum_k uv_{jgt}^{*k} q_{jgt}^k$ denote

the export shares for country j , the Laspeyres index in (D4) equals $P_{ijgt}^L = \sum_k s_{jgt}^{*k} (uv_{igt}^{*k} / uv_{jgt}^{*k})$.

Likewise, the Paasche index is a weighted average of the unit-value ratios using the export shares

s_{igt}^{*k} of country i . In either case, we can alternatively use the ratio of quality-adjusted unit values,

$(UV_{igt}^{*k} / UV_{jgt}^{*k})$, as defined in (D3). 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,

$P_{ijgt}^F = (P_{ijgt}^L P_{ijgt}^A)^{0.5}$. Then the GEKS price index of country i relative to k is computed by taking

the mean over all Fisher indexes for exports of country i relative to exports of j times the Fisher index for exports of j relative to exports of k :

$$P_{ikgt}^{GEKS} \equiv \prod_{j=1}^C \left(P_{ijgt}^F P_{jkgt}^F \right)^{1/C}, \quad (D5)$$

with $P_{iigt}^F \equiv 1$ for $i=1, \dots, C$. In most applications, the resulting GEKS indexes are transitive.⁴² That property does not necessarily hold in our case, however, because two countries may not export the 4-digit SITC product to the same set of partners, so that the mean in (D5) 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 GEKS transformation of the Fisher Ideal indexes in (D5) 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.⁴³

This GEKS aggregation is done for each 4-digit SITC product. We trim one percent of the estimated quality-adjusted price indexes (i.e. the upper and lower 0.5 percent) and then proceed with the second stage aggregation over the SITC products g . We again use Fisher Ideal indexes – now computed by summing over products rather than over partner countries as in (D4) – together with the GEKS transformation. In this second step we choose the United States as the comparison country k , so we end up with indexes of unit values, or quality-adjusted unit-values, for each exporting country and year relative to the United States. These indexes are computed for all exports and for the one-digit Broad Economic Categories (BEC), distinguishing food and beverages, other consumer goods, capital, fuels, intermediate inputs, and transport equipment, so this breakdown should be useful for other researchers interested in international prices.

⁴² This is shown from (C1) by noting that $P_{jkgt}^{GEKS} = 1 / P_{kjgt}^{GEKS}$, so that we readily compute $P_{ikgt}^{GEKS} P_{kmgt}^{GEKS} = P_{imgt}^{GEKS}$.

⁴³ 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.

Our treatment of imports is similar to our treatment of exports, so we only highlight the differences. In the first stage, the Laspeyres and Paasche indexes are computed by summing over source countries i that importers k and l both purchase from. So we compare the import prices of countries k and l only if they come from the same exporter i . As we found earlier, the Laspeyres and Paasche indexes can be expressed as share-weighted averages of the net-of-tariff c.i.f. unit-value ratio, or quality-adjusted unit-value ratio as in (D3), for countries k relative to l . That is, the Laspeyres and Paasche indexes depend on uv_{igt}^k / uv_{igt}^l , or alternatively on the quality-adjusted unit value UV_{igt}^k / UV_{igt}^l . We then compute the Fisher Ideal indexes and perform the GEKS transformation, resulting in an index of the import prices for country k relative to a base country m for each SITC product.⁴⁴ In the second stage, we aggregate over products g to obtain indexes of import prices, and quality-adjusted prices, relative to the United States for each BEC category. Dividing the former by the latter, we obtain the import index of quality.

Appendix E: Estimation

For convenience, we omit the goods subscript g in what follows, though all parameters and equations differ by SITC good. To utilize the GMM methodology introduced by Feenstra (1994), we need to develop the supply side in more detail. The c.i.f. and f.o.b. prices shown in (8) depend on the iceberg and the specific transport costs. The former depends on one plus the *ad valorem* tariffs, denoted by tar_{it}^k , and we model both costs as also depending log-linearly on the distance from country i to k and the aggregate physical export quantity (X_{it}^k / uv_{it}^k):

$$\ln \tau_{it}^k = \eta_t + \eta_0 \ln tar_{it}^k + \eta_1 \ln dist_i^k + \eta_2 \ln(X_{it}^k / uv_{it}^k) + \xi_{lit}^k, \quad (E1)$$

⁴⁴ Analogous to the export side, for each 4-digit SITC product and year we chose the base country l as the importer having the largest number of source countries times its total imports from all of them.

$$\ln T_{it}^k = \chi_t + \chi_1 \ln dist_i^k + \chi_2 \ln(X_{it}^k / uv_{it}^k) + \xi_{2it}^k. \quad (E2)$$

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

Notice from (7a) and (7b) we can write:

$$\left[\left(\ln \overline{p_{it}^k} - \ln \overline{p_{jt}^k} \right) - \theta \left(\ln \overline{p_{it}^{*k}} - \ln \overline{p_{jt}^{*k}} \right) \right] = \ln \tau_{it}^k + (1 - \theta) \ln T_{it}^k. \quad (E3)$$

Substituting for prices in (E3) using (20) and for trade costs using (E1) and (E2), we write an inverse supply curve using a similar linear combination of c.i.f. and f.o.b. unit values that appear in the demand equation (22):

$$\begin{aligned} \left[\left(\ln uv_{it}^k - \ln uv_{jt}^k \right) - \theta \left(\ln uv_{it}^{*k} - \ln uv_{jt}^{*k} \right) \right] &= (\eta_0 - 1) (\ln tar_{it}^k - \ln tar_{jt}^k) \\ &+ \omega_1 (\ln dist_i^k - \ln dist_j^k) + \omega_2 \left[\ln(X_{it}^k / uv_{it}^k) - \ln(X_{jt}^k / uv_{jt}^k) \right] + (\tilde{\xi}_{it}^k - \tilde{\xi}_{jt}^k), \end{aligned} \quad (E4)$$

where $\omega_i \equiv (\eta_i + (1 - \theta)\chi_i)$, $i=1,2$, and $\tilde{\xi}_{it}^k \equiv \xi_{1it}^k + u_{it}^k + (1 - \theta)\xi_{2it}^k - \theta u_{it}^{*k}$ incorporates the measurement error in (20). We rewrite (E4) slightly by shifting the export values and unit values to the left:

$$\begin{aligned} \left[(1 + \omega_2) \left(\ln uv_{it}^k - \ln uv_{jt}^k \right) - \theta \left(\ln uv_{it}^{*k} - \ln uv_{jt}^{*k} \right) \right] &- \omega_2 (\ln X_{it}^k - \ln X_{jt}^k) \\ &= (\eta_0 - 1) (\ln tar_{it}^k - \ln tar_{jt}^k) + \omega_1 (\ln dist_i^k - \ln dist_j^k) + (\tilde{\xi}_{it}^k - \tilde{\xi}_{jt}^k). \end{aligned} \quad (E5)$$

We combine this supply curve with the demand equation (22), rewritten as:

$$\begin{aligned} \ln X_{it}^k - \ln X_{jt}^k + A^k \left[\left(\ln uv_{it}^k - \ln uv_{jt}^k \right) - \alpha^k \theta \left(\ln uv_{it}^{*k} - \ln uv_{jt}^{*k} \right) \right] &= \delta_0 (\ln L_{it} - \ln L_{jt}) \\ &+ \delta_i - \delta_j - B^k \beta'(F_i^k - F_j^k) - C^k \left(\ln tar_{it}^k - \ln tar_{jt}^k \right) + \tilde{\varepsilon}_{it}^k - \tilde{\varepsilon}_{jt}^k, \end{aligned} \quad (E6)$$

where $\tilde{\varepsilon}_{it}^k \equiv (\varepsilon_{it}^k + A^k u_{it}^k - A^k \alpha^k \theta u_{it}^{*k})$ includes the measurement error in (20), and $C^k \equiv A^k + B^k$.

For convenience, we have grouped together the tariff terms and will treat these as control variables below.,

Taking the product of (E5) and (E6) and dividing by $A^k(1 + \omega_2)$, we obtain:

$$\begin{aligned}
& \left(\ln uv_{it}^k - \ln uv_{jt}^k \right)^2 \\
&= \left[\alpha^k \theta + \frac{\theta}{(1 + \omega_2)} \right] \left(\ln uv_{it}^k - \ln uv_{jt}^k \right) \left(\ln uv_{it}^{*k} - \ln uv_{jt}^{*k} \right) - \frac{\alpha^k \theta^2}{(1 + \omega_2)} \left(\ln uv_{it}^{*k} - \ln uv_{jt}^{*k} \right)^2 \\
&+ \frac{\omega_2}{A^k(1 + \omega_2)} \left(\ln X_{it}^k - \ln X_{jt}^k \right)^2 + \left(\frac{\omega_2}{(1 + \omega_2)} - \frac{1}{A^k} \right) \left(\ln X_{it}^k - \ln X_{jt}^k \right) \left(\ln uv_{it}^k - \ln uv_{jt}^k \right) \\
&+ \left(\frac{\theta}{A^k(1 + \omega_2)} - \frac{\alpha^k \omega_2 \theta}{(1 + \omega_2)} \right) \left(\ln X_{it}^k - \ln X_{jt}^k \right) \left(\ln uv_{it}^{*k} - \ln uv_{jt}^{*k} \right) + Controls_{it}^k + \mu_{it}^k,
\end{aligned} \tag{E7}$$

with the control terms,

$$\begin{aligned}
Controls_{it}^k &\equiv \frac{1}{A^k(1 + \omega_2)} \left[(\eta_0 - 1) (\ln tar_{it}^k - \ln tar_{jt}^k) + \omega_1 (\ln dist_i^k - \ln dist_j^k) \right] \times \\
&\left[\delta_0 (\ln L_{it} - \ln L_{jt}) + \delta_i - \delta_j - B^k \beta' (F_i^k - F_j^k) - C^k (\ln tar_{it}^k - \ln tar_{jt}^k) \right]
\end{aligned} \tag{E8}$$

and the error term,

$$\begin{aligned}
\mu_{it}^k &= \frac{(\tilde{\xi}_{it}^k - \tilde{\xi}_{jt}^k)}{A^k(1 + \omega_2)} \left[\delta_i - \delta_j + \delta (\ln L_{it} - \ln L_{jt}) - B^k \beta' (F_i^k - F_j^k) + \tilde{\varepsilon}_{it}^k - \tilde{\varepsilon}_{jt}^k \right] \\
&+ \frac{(\tilde{\varepsilon}_{it}^k - \tilde{\varepsilon}_{jt}^k)}{A^k(1 + \omega_2)} \left[(\eta_0 - 1) (\ln tar_{it}^k - \ln tar_{jt}^k) + \omega_1 (\ln dist_i^k - \ln dist_j^k) \right].
\end{aligned}$$

We treat the country fixed effects, sectoral labor force, distance, tariffs and language variables for the fixed costs of exporting as exogenous, so they are uncorrelated with the demand and supply shocks. We further assume that the supply and gravity shocks are uncorrelated in (E5) and (E6), so that $E\mu_{it}^k = 0$ for each source country i and destination k . This is the moment condition that we use to estimate (E7). This equation is simplified using $\zeta^k = \gamma / [\alpha^k \theta (\sigma - 1)]$ and so $\zeta^k \alpha^k = \gamma / [\theta (\sigma - 1)] = \zeta^{US} \alpha^{US} = \zeta^{US}$, since $\alpha^{US} \equiv 1$ by normalization. It follows that

$\gamma = \zeta^{US} \theta(\sigma - 1)$, and then for (E6) we obtain $A^k = (\sigma - 1)[1 + \zeta^{US} \theta(\sigma - 1)] / [1 + \alpha^k \theta(\sigma - 1)]$ and $B^k = [\zeta^{US} \theta(\sigma - 1) - \alpha^k \theta(\sigma - 1)] / [1 + \alpha^k \theta(\sigma - 1)]$. Substituting these relations into (E7) and (E8), we obtain an equation that is nonlinear in the parameters θ and σ .

For estimation, we average the variables in (E7) and (E8) over time, which eliminates the time subscript and gives a cross-country regression that can be estimated with nonlinear least squares (NLS). Another challenge is to incorporate the source country fixed effects ($\delta_i - \delta_j$) interacted with distance and tariffs as appear in (E8). 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 (E7) on the source country fixed effects and their interaction terms, and then estimate (E7) using the residuals obtained from these preliminary regressions. The source country fixed effects are needed to control for the measurement errors in the c.i.f. and f.o.b. unit values, shown in (20), which we assume are independent of each other and of the export values. Then the variance of the measurement errors appears in the error term after averaging over time, and the source country fixed effects absorb these variances.

A final challenge is to estimate the destination country's preference for quality, α^k . Equation (7a) provides us with a method to estimate these preferences using data on f.o.b. unit values, which we assumed to be linked to f.o.b. prices in (20) by $\ln uv_i^{*k} = \ln \overline{p_i^{*k}} + u_i^{*k}$, with measurement error u_i^{*k} . We model α^k as depending on real GDP per capita of country k from the Penn World Table. Taking logs of (7a), adding a time subscript t and a SITC goods subscript g , and assuming that specific transport costs depend on distance, we estimate:

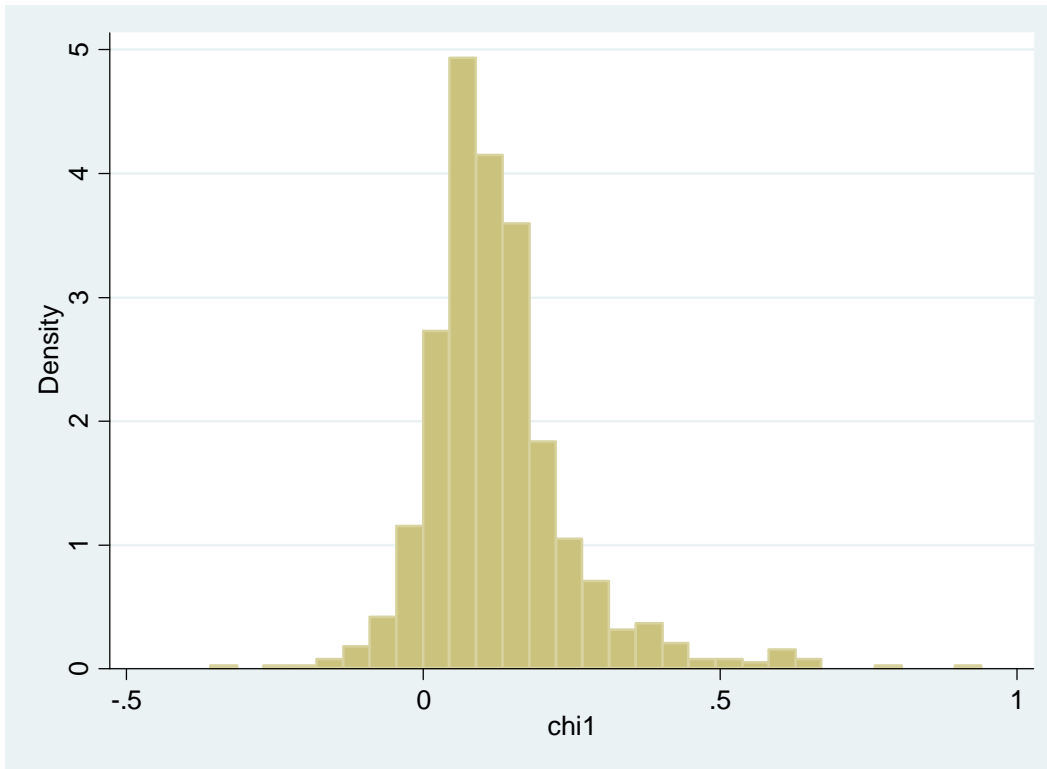
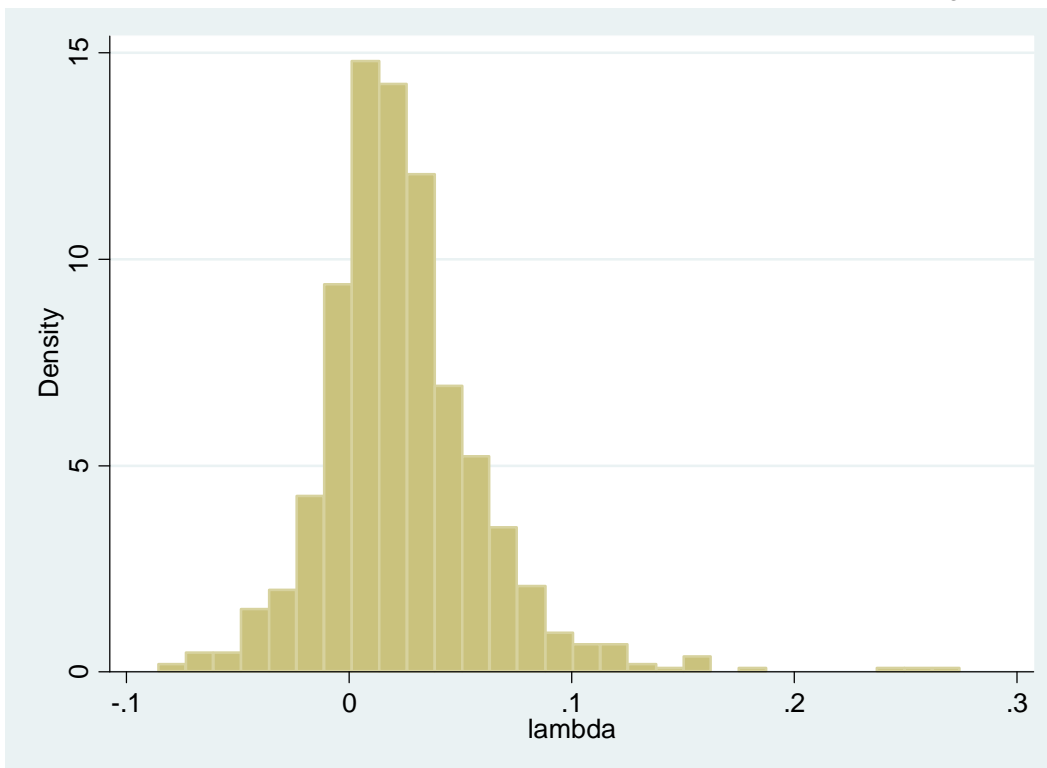
$$\ln uv_{it}^{*k} = \chi_{it} + \chi_1 \ln \text{dist}_i^k + \ln \left[\left(\frac{1}{1 - \alpha_i^k \theta} \right) \left(\frac{\sigma}{\sigma - 1} \right) - 1 \right] + u_{it}^{*k}, \quad (\text{E9})$$

with,
$$\alpha_t^k = 1 + \lambda \ln \left(\text{RGDPL}_t^k / \text{RGDPL}_t^{US} \right). \quad (\text{E10})$$

(E9) is consistent with the specific transport costs in (E2), except that we now ignore congestion ($\chi_2 = 0$) and generalize to allow for a source-country-time fixed effect χ_{it} . We measure real GDP per capita, RGDPL_t^k , relative to that in the United States as a normalization. Substituting (E10) into (E9), this nonlinear regression is run for each SITC 4-digit industry over 1984-2011. We began with initial estimates of σ and θ obtained from the GMM system (E7)-(E8) estimated with $\alpha_t^k \equiv 1$, obtaining initial estimates of λ and therefore α_t^k using nonlinear least squares. The average over time of these new values α_t^k are substituted into (E7)-(E8), and we re-estimate the GMM system to obtain new estimates for σ and θ . We iterated on this procedure several times, and found that the distribution of estimates of σ and θ quickly converged.

In Appendix Figure 3 we show the frequency of estimates for χ_1 , the coefficient on log distance. Its median value over 862 4-digit SITC industries is 0.10. Over 85 percent of estimates are significantly positive, while under 6 percent of the estimates are significantly negative. The fact that the f.o.b. unit value – which is *net* of transport costs – is increasing in distance is interpreted by Hummels and Skiba (2004) as evidence of the “Washington apples” effect, whereby quality grows with distance.

In Appendix Figure 4 we show the frequency distribution for estimates of λ , the coefficient of real GDP per capita in determining α_t^k . Its median value over the 4-digit SITC industries is 0.021, over 70 percent of estimates are significantly positive and 14% of estimates are significantly negative. The negative estimates can be explained by plausible cases where lower-income countries prefer higher quality due to the changing composition of goods within SITC 4-digit categories. A leading example is SITC 3341, “Gasoline and other Light Fuels,” which includes fuels for aircraft. It has $\lambda = -0.06$, one of the largest significant negative values,

Appendix Figure 3: Frequency Distribution for Estimates of χ_{1g} **Appendix Figure 4: Frequency Distribution for Estimates of λ_g** 

since many small, low-income economies (especially island countries) without refining capacity require relatively more of the higher-quality aircraft fuel. We therefore retain the negative estimates of λ . The implied values for α_t^k range between 0.42 and 1.31 over all goods and countries (recalling that it is normalized to unity for the United States).

Because the GMM estimation is performed after eliminating the source country fixed effects and their interactions, we do not obtain the coefficients of those terms. Likewise, we do not recover the estimates of the other control terms in (E7). So a second-stage estimation is performed to obtain these coefficients, working from the gravity equation (B11). We substitute the estimates $\hat{\sigma}$, $\hat{\theta}$, $\hat{\gamma}$ and $\hat{\lambda}$ into (23) to obtain the coefficients \hat{A}^k , \hat{B}^k which appear in (B11). Substituting the net-of-tariff unit values from (20) into the prices in (B11), the coefficient of $\ln tar_i^k$ there becomes $\hat{C}^k \equiv \hat{A}^k + \hat{B}^k$. Then using $\ln[M_{it}(w_{it}/\phi_{it})^{-\gamma}] = \delta_i + \delta_0 \ln L_{it} + \varepsilon_{it}^k$, the gravity equation (B11) is run over time for each SITC good:

$$\ln X_{it}^k + \hat{A}^k \left(\ln uv_{it}^k - \hat{\alpha}^k \hat{\theta} \ln uv_{it}^{*k} \right) + \hat{C}^k \ln tar_{it}^k = \delta_{it}^k + \delta_i + \delta_0 \ln L_{it} - \hat{B}^k \beta' F_i^k + \hat{\varepsilon}_{it}^k, \quad (E11)$$

where the error term $\hat{\varepsilon}_{it}^k$ includes $\tilde{\varepsilon}_{it}^k$ plus the sampling error in the coefficients $\hat{\sigma}$, $\hat{\theta}$, $\hat{\gamma}$ and $\hat{\lambda}$, and δ_{it}^k incorporates the log of (Y_t^k / R_t^k) from (B11), as well as κ_{it}^k and $(Y_t^k / p_t^k)^{\beta_0}$ from \tilde{f}_{it}^k . Running (E11) as a fixed-effects regression for each SITC good and each year, we obtain the coefficients $\hat{\beta}$ from which we construct the fixed costs of exporting $\exp(\hat{\beta}' F_i^k)$ that enter into the quality-adjusted price calculations.