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ESTIMATING CROSS-COUNTRY DIFFERENCES IN PRODUCT QUALITY

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**ABSTRACT**

We develop a method for decomposing countries' observed export prices into quality versus quality-adjusted-price components using information contained in their trade balances. Holding observed export prices constant, countries with surpluses are inferred to offer higher quality than countries running deficits. Our method accounts for variation in trade balances induced by horizontal and vertical differentiation. We use our method to examine manufacturing product quality among the world's top exporters from 1989 to 2003. We find that the initial quality gap between high- and low-income countries is smaller than their initial income gap, and that the former narrows considerably faster over time.

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## 1. Introduction

Theoretical and empirical research increasingly point to the importance of product quality in international trade and economic development.<sup>1</sup> Unfortunately, relatively little is known about how countries' product quality varies across time, or how it is influenced by trade liberalization and other aspects of globalization. A major impediment to research in this area is lack of data – reliable estimates of export quality for a wide range of countries, industries and years do not exist. In this paper, we introduce a method for obtaining such estimates that incorporates information about world demand for countries' products.

Researchers often react to the absence of information about countries' product quality by constructing *ad hoc* proxies, the most common of which is observed export prices (unit values).<sup>2</sup> This measure is unsatisfactory, however, because export prices may vary for reasons other than quality. Chinese shirts might be cheaper than Italian shirts in the U.S. market because of lower quality, but they might also sell at a discount because China has lower production costs or an undervalued exchange rate. If consumers value variety and goods are horizontally as well as vertically differentiated, high-cost producers can survive in the U.S. market even in the face of cost disadvantages.

Our method for identifying countries' export quality involves decomposing observed export prices into quality versus quality-adjusted-price components. We define quality to be any tangible or intangible attribute of a good that increases all consumers' valuation of it. Countries' product quality relative to a numeraire country is identified by combining data on their observed export prices with information about global demand for their products contained in their trade balance *vis a vis* the world. The intuition behind our identification is straightforward and has been used extensively in the industrial organization literature: because consumers are assumed to care about price relative to quality in choosing among products, two countries with the same export prices but different global trade balances must have products with different levels of quality. Among countries with identical export prices, the country with the higher trade balance is revealed to possess

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<sup>1</sup>Flam and Helpman (1987) is representative of a line of theoretical research studying the influence of product quality on international trade. Empirically, cross-country and time-series variation in product quality has been linked to firms' export success (Brooks 2006, Verhoogen 2007), countries' skill premia (Verhoogen 2007), quantitative import restrictions (Aw and Roberts 1986, Feenstra 1988) and trade patterns (Schott 2004, Hallak 2006). The contribution of quality growth to macroeconomic growth is investigated theoretically by Grossman and Helpman (1991) and empirically by Hummels and Klenow (2005).

<sup>2</sup>See, for example, Schott (2008). More generally, unit value differences figure prominently in surveys of countries' "quality competitiveness" (e.g., Aiginger 1998, Verma 2002, Ianchovichina et al. 2003, and Fabrizio et al. 2007) and also are often used to distinguish horizontal from vertical intra-industry trade flows (e.g., Abed-el-Rahman 1991 and Aiginger 1997).

higher product quality.<sup>3</sup>

A major contribution of the paper is to generalize this intuition to a setting where countries are also allowed to differ in the number of *unobserved* horizontal varieties they export in each product category (e.g., red versus blue men’s long-sleeve cotton shirts), a standard assumption in trade models. Accounting for unobserved horizontal differentiation is difficult because it introduces an additional factor besides quality that can increase consumer demand for a country’s products. All else equal, consumer love of variety implies that countries producing a larger number of varieties in a product category export larger quantities and therefore exhibit higher trade surpluses. Unless the number of horizontal varieties that countries export is accounted for, this increase in net trade will be interpreted, erroneously, as higher product quality. Our approach assumes a negative relationship between quality-adjusted prices and the number of varieties countries export. We justify this assumption by appealing to theoretical findings in Romalis (2004) and Bernard et al. (2007) that demonstrate that countries’ comparative advantage sectors exhibit both relatively low prices – due to relatively low factor costs – and a relatively high number of varieties – due to disproportionate use of factor inputs.<sup>4</sup>

Using countries’ net trade with the rest of the world to identify consumer demand imposes an important practical constraint on empirical implementation of our method. Currently, the most reliable time-series data on countries’ trade balances are recorded according to relatively coarse industries relative to the much more disaggregate products (e.g., men’s long-sleeve cotton shirts) at which some countries’ export prices can be observed. To deal with this constraint we derive a theoretically appropriate price index that aggregates countries’ product-level export prices up to the industry level. We refer to this index as the “Impure Price Index”. We term it “impure” because its prices are “contaminated” by quality. This index has the useful property of being separable into quality versus quality-adjusted-price components. It is developed under the assumption that countries’ export quality is constant across products within industries. This assumption creates an “aggregation trade-off”: while product quality is more likely to be constant across products the more disaggregate the industry, data on countries’ global net trade

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<sup>3</sup>The use of market shares to infer unobserved consumer valuation is well-established in the industrial organization and index number literatures (e.g. Berry 1994 and Bils 2004, respectively). Here, countries’ net trade with the rest of the world (conditional on trade costs) is a natural expression of their “market share”.

<sup>4</sup>Feenstra (1994) outlines a method for computing import price indexes that accounts for the introduction of new product varieties (see also Broda and Weinstein 2006). Given its focus on changes in prices over time, that methodology requires no knowledge of cross-sectional variation in the number of varieties countries export within product categories so long as that number is constant over time for a subset of countries. Here, we allow the number of varieties to vary across all countries.

becomes more scarce as well as more susceptible to measurement error. Use of disaggregate industries may also be problematic if countries' use of intermediate inputs straddles the industries at which quality is being estimated. Separate estimations of quality for apparel and textiles, for example, might not capture the fact that some countries import textiles to produce apparel. As a result, textile and apparel quality might be under- and over-estimated in these countries.

We show that even though the Impure Price Index comparing two countries' export prices is unobservable, it is bounded by observable Paasche and Laspeyres indexes defined over their common exports to a third country (e.g., the United States). This result anchors a two-stage strategy for inferring countries' export quality. In the first stage, we use the large set of bilateral Paasche and Laspeyres bounds (e.g., Germany versus China, Switzerland versus Germany, France versus Thailand, etc.) to estimate the Impure Price Index for each country-industry-year relative to a common numeraire. In the second stage, we use data on countries' global net trade in the industry to strip away variation in quality-adjusted (or "pure") prices from the estimated Impure Price Indexes. This procedure yields estimates of quality that vary by country, industry and year.

We use our method to estimate manufacturing quality for the world's 43 largest exporters over the period 1989 to 2003. The estimated Quality Indexes reveal substantial variation in quality levels across countries in any given year as well as across years. We find that relative export quality for overall manufacturing increases most dramatically for Ireland, Malaysia and Singapore over the sample period, and falls most dramatically for Hong Kong, New Zealand and Australia. Among countries that begin the sample period in the top tercile of quality, Japan and Spain experience the largest (relative) declines. We also show that our estimates of export quality and their evolution over time can deviate substantially from estimates of export quality based on raw export prices. Indeed, we find that our Quality Indexes and raw export prices move in opposite directions for one third of the countries in our sample, including some of those with the largest increases in our quality estimates. Finally, dividing our sample in halves according to income per capita in 1989 (high versus low), we find that the mean quality gap between the two groups is initially smaller than the gap in mean per capita income, and that the mean quality gap narrows considerably faster.

This paper's focus on cross-sectional variation in product quality differentiates it from a very large index number literature devoted to constructing quality-adjusted cost-of-living indexes. Here, rather than measure quality changes in bundles of products purchased over time, we identify quality variation over simultaneously purchased bundles from different

sources of supply. Since we cannot observe products' underlying attributes, we are also unable to make use of standard strategies – such as hedonic pricing – that link product attributes to specific dimensions of quality.<sup>5</sup> Our method complements such efforts, however, because its use of publicly available trade data permits estimation of product quality across a broad range of countries, industries and years for which surveys of product characteristics may be unavailable or prohibitively expensive to collect.<sup>6</sup>

Our analysis is more closely related to previous attempts in the international trade literature to deal with potential variation in unit values not entirely due to variation in product quality. Hallak (2006), for example, assumes a monotonic relationship between per-capita income and “pure prices” at the sector level while, in the closest precedent to this paper, Hummels and Klenow (2005) use import prices and quantities to make inferences about the cross-sectional elasticity of quality with respect to country income and size. Neither of these papers, however, permits explicit estimation of product quality by country, sector, and year, as is done in this paper.<sup>7</sup> Our approach is also different from an earlier strand of literature primarily interested in analyzing the effect of import quotas on the quality composition of trade (e.g. Aw and Roberts 1986, Boorstein and Feenstra 1987, and Feenstra 1988). In this literature, import quality increases when the composition of imports shifts toward high-quality product categories. This literature therefore adopts an across-product-category view of quality variation and quality upgrading in contrast to the within-product-category view pursued here.

Estimates of countries' export quality will obviously find use in testing models of international specialization and development. They may also benefit other fields, such as productivity and growth, where, despite the existence of an influential theoretical literature linking the production of quality to economic growth (e.g., Grossman and Helpman 1991, Aghion and Howitt 1992), empirical investigation of this link is scarce due to the unavailability of reliable measures of quality. Controlling for product quality is also crucial for computing the import and export price indexes used to deflate national accounts aggregates. Current estimates of “real GDP” in the Penn World Tables, for example, deflate nominal GDP using a purchasing-power-parity index based on final expenditure data. As noted by Feenstra et al. (2007), this adjustment likely does not appropriately

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<sup>5</sup>Feenstra (1995), for example, demonstrates how information on product attributes can be used to establish bounds on the exact hedonic price index.

<sup>6</sup>The International Price Program of the U.S. Bureau of Labor Statistics constructs import and export price indexes by combining survey data on firms' prices with firms' assessments about changes in the quality of their products over time (Alterman et al. 1999).

<sup>7</sup>More recently, Khandelwal (2007) has developed a method for estimating quality based on the assumption of a nested logit demand system.

capture changes in countries' *production* over time because it ignores the terms of trade. The ability to net quality out of countries' import and export price indexes before performing the terms-of-trade correction would enhance the reporting of national accounts. Development of country-sector specific quality-adjusted price indexes is also likely to be useful in the labor economics literature. The distributional consequences of international trade implied by the Stolper-Samuelson theorem, for example, cannot be properly identified if the import and export price changes used to empirically assess the theorem's relevance do not properly account for changes in countries' product quality.<sup>8</sup>

The paper is structured as follows. Section 2 outlines our assumptions about consumer demand and introduces the Impure and Pure Price indexes that will be the focus of our analysis. Section 3 shows that the unobservable Impure Price Index is bounded by observable Paasche and Laspeyres indexes. Section 4 derives the relationship between the Pure Price Index and countries' sectoral net trade. Sections 5 through 7 describe the application of our method to identifying export quality trends for 43 large trading countries over the period 1989 to 2003. Section 8 concludes.

## 2. Preferences and Price Indexes

This section describes the preference structure underlying our analysis and formally introduces the price and quality indexes that are the focus of our method.

### 2.1. Preferences

Goods are classified into product categories, which are in turn classified into sectors. Sectors are indexed by  $s = 1, \dots, S$ , while product categories (within sectors) are indexed by  $z = 1, \dots, Z_s$ . In our empirical investigation below, product categories correspond to ten-digit Harmonized System (HS) categories, the finest possible level of aggregation in trade statistics.

There are  $C$  countries, indexed by  $c = 1, \dots, C$ . The theoretical framework presented below focuses on one sector,  $s$ . The analysis for other sectors to which the method is applied is analogous.

Preferences are common across countries, and are represented by a two-tier utility function that incorporates consumer love of variety. The upper tier is Cobb-Douglas, with

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<sup>8</sup>See Goldberg and Pavcnik (2007) for a discussion of the empirical validity of the Stolper-Samuelson mechanism.

expenditure shares  $b_s$  for each sector  $s$ . The lower tier has the following CES form<sup>9,10</sup>

$$u_s = \left[ \sum_c \sum_z (\xi_z \lambda_s^c x_z^c)^{\frac{\sigma_s-1}{\sigma_s}} n_z^c \right]^{\frac{\sigma_s}{\sigma_s-1}}, \quad \sigma_s > 1. \quad (1)$$

In the sub-utility function (1),  $n_z^c$  is the number of horizontally differentiated varieties of product  $z$  produced by country  $c$ ,  $x_z^c$  is the quantity consumed per variety, and  $\sigma_s$  is the elasticity of substitution between varieties. We note that by indexing product categories instead of varieties, we implicitly assume symmetry across varieties in the same product category. The utility function includes two shifters,  $\xi_z$  and  $\lambda_s^c$ . The first shifter,  $\xi_z$ , varies across product categories but is constant across countries for a particular product category. It captures consumers' valuation of the essential characteristics common to the heterogeneous varieties of a particular product category. Consumers, for example, might have a higher preference for table varieties than chair varieties. The second shifter,  $\lambda_s^c$ , varies across countries and sectors, but is constant across products within a particular country and sector. It represents product quality and captures the combined effect of all product characteristics, other than price and those already captured by  $\xi_z$ , on consumers' valuation of a good. Product quality encompasses both physical attributes (e.g., durability) and intangible attributes (e.g., product image due to advertising). These assumptions are formalized as follows:

**Assumption 1:**  $\xi_z^c = \xi_z, \quad \forall c = 1, \dots, C, \quad \forall z = 1, \dots, Z_s.$

**Assumption 2:**  $\lambda_z^c = \lambda_s^c, \quad \forall z = 1, \dots, Z_s.$

With the preference structure defined by equation (1), product demand depends on quality-adjusted or “pure” prices. Letting  $p_z^c$  be the export price of a typical variety of product  $z$  produced in country  $c$ , we define the “pure” price of that variety by  $\tilde{p}_z^c = p_z^c / (\xi_z \lambda_s^c)$ . The pure price is a quality-adjusted price. It is also divided here by  $\xi_z$  for notational compactness, but none of the results or their interpretation is affected by this choice.

<sup>9</sup>The assumption of homothetic preferences, although standard in the international trade literature, is potentially strong in this context. It is worthy of more focused attention in future theoretical and empirical research.

<sup>10</sup>To simplify notation, subindexes on summations refer to all members of a set unless otherwise noted, e.g.  $\sum_c$  and  $\sum_{c'}$  both sum over all countries  $c = 1, \dots, C$  while  $\sum_{c' \neq c}$  sums over all countries except  $c$ . For product categories,  $\sum_z$  denotes the sum across all product categories in sector  $s$ ,  $z = 1, \dots, Z_s$ . Note that we omit subindex  $s$  from  $z$ .

## 2.2. The Pure and Impure Price Indexes

In this section we develop notation to keep track of countries' unobserved numbers of varieties. Define  $\bar{n}_s^c$  to be the average number of varieties across product categories produced by country  $c$  in sector  $s$ ,

$$\bar{n}_s^c = \frac{1}{Z_s} \sum_z n_z^c \quad (2)$$

and define  $\bar{n}_z$  to be the (country  $o$ -normalized) world average number of varieties of product  $z$ ,

$$\bar{n}_z = \frac{1}{C} \sum_c n_z^c \frac{\bar{n}_s^o}{\bar{n}_s^c} \quad (3)$$

The normalization in (3) re-scales the number of varieties of each country into common, country- $o$  units, according to the ratio of the average number of varieties between  $o$  and  $c$ . Define also  $\tilde{n}_z^c$  to be country  $c$ 's "excess variety" in product  $z$  relative to the world average,

$$\tilde{n}_z^c = n_z^c \frac{\bar{n}_s^o}{\bar{n}_s^c} - \bar{n}_z. \quad (4)$$

Note that excess variety has the convenient property that  $\sum_z \tilde{n}_z^c = 0, \forall c = 1, \dots, C$ .

Define an aggregator<sup>11</sup> of product prices produced in country  $c$  and sector  $s$  as

$$P_s^c = \left[ \sum_z \bar{n}_z \xi_z^{\sigma_s - 1} (p_z^c)^{1 - \sigma_s} \right]^{\frac{1}{1 - \sigma_s}}. \quad (5)$$

The Impure Price Index between countries  $c$  and  $d$  is then defined as

$$P_s^{cd} = P_s^c / P_s^d. \quad (6)$$

The Impure Price Index is a summary measure of price variation between goods produced by countries  $c$  and  $d$  in sector  $s$ . The index is "impure" in the sense that it is defined over prices that are "contaminated" by quality. The index is transitive, so that  $P_s^{cd} P_s^{do} = P_s^{co}$ . Choosing country  $o$  as the numeraire country, we can associate an index number,  $P_s^{co}$ , with each country  $c$ , noting that  $P_s^{cd}$  can always be recovered from the ratio  $P_s^{co} / P_s^{do}$ . In particular, the value of this ratio is independent of which country is chosen as the numeraire.

<sup>11</sup>This type of price aggregator is often called a price "index" in the trade literature (e.g., Anderson and van Wincoop 2004). We reserve the term "index" here for price comparisons between countries, in accordance with terminology employed in the index number literature.

The Impure Price Index can be decomposed into an index of quality and an index of pure prices:

$$P_s^{cd} = \tilde{P}_s^{cd} \lambda_s^{cd}, \quad \lambda_s^{cd} = \frac{\lambda_s^c}{\lambda_s^d}, \quad \tilde{P}_s^{cd} = \frac{\tilde{P}_s^c}{\tilde{P}_s^d} = \left[ \frac{\sum_z \bar{n}_z (\tilde{p}_z^c)^{1-\sigma_s}}{\sum_z \bar{n}_z (\tilde{p}_z^d)^{1-\sigma_s}} \right]^{\frac{1}{1-\sigma_s}} \quad (7)$$

The Quality Index,  $\lambda_s^{cd}$ , between countries  $c$  and  $d$  in sector  $s$  is simply defined as the ratio of the two countries' quality levels. The Impure Price Index and the Quality Index implicitly define the Pure Price Index,  $\tilde{P}_s^{cd}$ . The Pure Price Index is a summary measure of pure price variation between countries, and it is also transitive. Combining estimates of countries' sectoral Impure Price Indexes with inferences about their sectoral Pure Price Indexes derived from their global net trade in the sector, we will use the decomposition in (7) to identify countries' sectoral relative product quality.

### 3. Bounding the “Impure” Price Index

The bilateral Impure Price Index in equation (7) cannot be observed because it depends upon unobservables such as the number of varieties exported by the country pair and the elasticity of substitution. In this section we outline a set of assumptions which allow the Impure Price Index to be bounded by observable Paasche and Laspeyres indexes defined over the two countries' common exports to a third country. In Section 5, we demonstrate how overlapping bilateral bounds across country pairs can be used to identify Impure Price Indexes for all countries (relative to a numeraire country).

#### 3.1. Constrained Expenditure Function

We focus on countries' exports to a single “common importer”, which we refer to as the United States given the focus of our empirical examination below. The analysis would be identical were it to be applied to any other common importer.

We define a country as “active” in product  $z$  if it reports positive exports to the United States in that category. Let  $I_s$  be the set of all product categories in sector  $s$ , and let  $I_s^c$  be the subset of active categories in country  $c$ . Define vector  $\mathbf{p}_s$  to include the U.S. import prices of all active categories in sector  $s$  from all countries. Define analogously vectors  $\mathbf{q}_s$ ,  $\mathbf{n}_s$ ,  $\boldsymbol{\lambda}_s$ , and  $\boldsymbol{\xi}_s$ . A vector of per-variety consumption  $\mathbf{x}_s$  is implicitly defined by  $\mathbf{q}_s$  and  $\mathbf{n}_s$ . Finally, stack these vectors across sectors to form vectors  $\mathbf{p}$ ,  $\mathbf{q}$ ,  $\mathbf{n}$ ,  $\boldsymbol{\lambda}$ ,  $\boldsymbol{\xi}$ , and  $\mathbf{x}$ .

Since our method is based on comparing import prices (as measured by unit values) across pairs of U.S. trading partners, we need to use notation specific to country pairs.

Index countries in a pair of U.S. trading partners by  $c$  and  $d$ . Denote by  $I_s^{cd}$  the set of active categories common to  $c$  and  $d$  in sector  $s$ .  $Z_s^{cd}$  is the number of such categories. Denote also by  $I_s^{c,-d}$  the set of products in which  $c$  is active but not  $d$ , by  $I_s^{d,-c}$  the set of products in which  $d$  is active but not  $c$ , and by  $U_s^{cd}$  the union of these two sets. Finally,  $\emptyset_s^{cd}$  is the set of products in which neither of the two countries is active. The set  $I_s$  can then be partitioned into  $I_s^{cd}$ ,  $U_s^{cd}$ , and  $\emptyset_s^{cd}$ .

We can use  $I_s^{cd}$  to break each of vectors  $\mathbf{p}$  and  $\mathbf{q}$  into two components. First, alternatively for each  $i = c, d$ ,  $\mathbf{p}_{s(cd)}^i$  and  $\mathbf{q}_{s(cd)}^i$  include prices and quantities, respectively, of exports by  $i$  of products in categories  $z \in I_s^{cd}$ . The remaining parts of  $\mathbf{p}$  and  $\mathbf{q}$  are denoted by  $\mathbf{p}_{s(cd)}^{-i}$  and  $\mathbf{q}_{s(cd)}^{-i}$ . These vectors include categories  $z \in I_s^{cd}$  exported by all countries other than  $i$ , and also categories  $z \notin I_s^{cd}$  exported by all countries (including  $i$ ).<sup>12</sup>

For a pair of exporting countries  $c$  and  $d$ , we now define the constrained expenditure (or import) function  $m_{s(cd)}^c(\mathbf{p}_{s(cd)}^i, \mathbf{q}_{s(cd)}^{-c}, \mathbf{n}, \boldsymbol{\lambda}, \boldsymbol{\xi}, U)$ . This function represents the minimum expenditure that the representative consumer in the U.S. would be required to spend on varieties exported by country  $c$  in categories  $z \in I_s^{cd}$  in order to attain utility level  $U$  when import prices of those varieties are  $\mathbf{p}_{s(cd)}^i$ , if this consumer is constrained to consume quantities  $\mathbf{q}_{s(cd)}^{-c}$  of all other products, and the number of varieties, quality, and product shifters are, respectively,  $\mathbf{n}, \boldsymbol{\lambda}, \boldsymbol{\xi}$ . The constrained expenditure function solves the problem

$$\min_{\mathbf{q}_{s(cd)}^c} \mathbf{p}_{s(cd)}^i \mathbf{q}_{s(cd)}^c \quad s.t. \quad U(\mathbf{q}_{s(cd)}^c, \mathbf{q}_{s(cd)}^{-c}, \mathbf{n}, \boldsymbol{\lambda}, \boldsymbol{\xi}) = U, \quad i = c, d \quad (8)$$

where  $U(\cdot)$  is the representative consumer utility function.<sup>13</sup>

By revealed preference, the minimum import expenditure on products produced by country  $c$  in categories  $z \in I_s^{cd}$ , when import prices of those products are  $\mathbf{p}_{s(cd)}^c$  while  $\mathbf{q}_{s(cd)}^{-c}, \mathbf{n}, \boldsymbol{\lambda}, \boldsymbol{\xi}$ , and  $U$  take their unconstrained equilibrium values, is the observed amount of imports:

$$m_{s(cd)}^c(\mathbf{p}_{s(cd)}^c, \mathbf{q}_{s(cd)}^{-c}, \mathbf{n}, \boldsymbol{\lambda}, \boldsymbol{\xi}, U) = \mathbf{p}_{s(cd)}^c \mathbf{q}_{s(cd)}^c. \quad (9)$$

However, when prices are  $\mathbf{p}_{s(cd)}^d$  instead of  $\mathbf{p}_{s(cd)}^c$ , the minimum import expenditure is equal to or lower than  $\mathbf{p}_{s(cd)}^d \mathbf{q}_{s(cd)}^c$ , because the amount  $\mathbf{p}_{s(cd)}^d \mathbf{q}_{s(cd)}^c$  is sufficient to attain utility  $U$  but  $\mathbf{q}_{s(cd)}^c$  is not necessarily optimal given  $\mathbf{p}_{s(cd)}^d$ . Hence,

$$m_{s(cd)}^c(\mathbf{p}_{s(cd)}^d, \mathbf{q}_{s(cd)}^{-c}, \mathbf{n}, \boldsymbol{\lambda}, \boldsymbol{\xi}, U) \leq \mathbf{p}_{s(cd)}^d \mathbf{q}_{s(cd)}^c. \quad (10)$$

<sup>12</sup>The term in parenthesis in the subindex denotes the subset of products within sector  $s$  in which countries  $c$  and  $d$  export in common to the U.S., i.e.  $\{z : z \in I_s^{cd}\}$ .

<sup>13</sup>Neary and Roberts (1980) and Anderson and Neary (1992) use the constrained expenditure function to analyze consumption choices under rationing.

Taking the ratio of (9) over (10), we obtain

$$M_{s(cd)}^c = \frac{m_{s(cd)}^c(\mathbf{p}_{s(cd)}^c, \mathbf{q}_{s(cd)}^{-c}, \mathbf{n}, \boldsymbol{\lambda}, \boldsymbol{\xi}, U)}{m_{s(cd)}^c(\mathbf{p}_{s(cd)}^d, \mathbf{q}_{s(cd)}^{-c}, \mathbf{n}, \boldsymbol{\lambda}, \boldsymbol{\xi}, U)} \geq \frac{\mathbf{p}_{s(cd)}^c \mathbf{q}_{s(cd)}^c}{\mathbf{p}_{s(cd)}^d \mathbf{q}_{s(cd)}^c} = H_s^{cd}. \quad (11)$$

Equation (11) displays a standard result in index number theory stating that the cost-of-utility price index  $M_{s(cd)}^c$  is larger than a Paasche price index,  $H_s^{cd}$ , defined here in a cross-sectional rather than a time-series context. The left-hand side of (11),  $M_{s(cd)}^c$ , captures the change in minimum expenditure on country  $c$ 's varieties (in categories  $z \in I_s^{cd}$ ) that would be necessary to maintain utility  $U$ , if import prices of those varieties changed from  $\mathbf{p}_{s(cd)}^d$  to  $\mathbf{p}_{s(cd)}^c$ , holding constant their number and characteristics (including quality), and the number, characteristics and quantity consumed of all other goods. The right-hand side of (11),  $H_s^{cd}$ , is a Paasche price index defined over the observed prices of the country pair's common exports to the U.S. in sector  $s$ .

Similarly, we can focus on imports from country  $d$  to obtain

$$M_{s(cd)}^d = \frac{m_{s(cd)}^d(\mathbf{p}_{s(cd)}^c, \mathbf{q}_{s(cd)}^{-d}, \mathbf{n}, \boldsymbol{\lambda}, \boldsymbol{\xi}, U)}{m_{s(cd)}^d(\mathbf{p}_{s(cd)}^d, \mathbf{q}_{s(cd)}^{-d}, \mathbf{n}, \boldsymbol{\lambda}, \boldsymbol{\xi}, U)} \leq \frac{\mathbf{p}_{s(cd)}^c \mathbf{q}_{s(cd)}^d}{\mathbf{p}_{s(cd)}^d \mathbf{q}_{s(cd)}^d} = L_s^{cd}, \quad (12)$$

where  $L_s^{cd}$  is a Laspeyres price index defined over the country pair's common exports to the U.S. in sector  $s$ . This is another standard result, which states that the cost-of-utility index  $M_{s(cd)}^d$  is bounded from above by a Laspeyres price index.<sup>14</sup>

We will now obtain explicit functional forms for  $M_{s(cd)}^c$  and  $M_{s(cd)}^d$ . Define  $\hat{u}_s \equiv \left[ \sum_{z \in I_s^{cd}} n_z^c (\xi_z \lambda_s^c x_z^c)^{\varphi_s} \right]^{1/\varphi_s}$ . The utility function  $U$  can be written as a function of  $\hat{u}_s$  and a function  $\bar{u}_s$  of arguments that are held constant in the minimization problem that defines the constrained expenditure function. Since  $U$  is strictly increasing in  $\hat{u}_s$ , there is a single value  $u_s^*$  of this variable that satisfies the constraint  $U(\hat{u}_s, \bar{u}_s) = U$  in (8). Then, we can rewrite the minimization problem as

$$\min_{x_z^c} \sum_{z \in I_s^{cd}} n_z^c p_z^i x_z^c \quad s.t. \quad \left[ \sum_{z \in I_s^{cd}} n_z^c (\xi_z \lambda_s^c x_z^c)^{\varphi_s} \right]^{1/\varphi_s} = u_s^*, \quad i = c, d.$$

The solution to this problem is the product between a CES aggregator measuring the unit

<sup>14</sup>Paasche and Laspeyres indexes are typically defined in a time series context, where there is a natural ordering of time periods. Since there is no natural ordering of countries in a multilateral context, calling one of these indexes Paasche and the other one Laspeyres rather than *vice versa* is arbitrary.

cost of utility and the target level of utility,  $u_s^*$ <sup>15</sup>

$$m_{s(cd)}^c(\mathbf{p}_s^i, \mathbf{q}_s^{-c}, \boldsymbol{\lambda}, \boldsymbol{\xi}, U) = \left[ \sum_{z \in I_s^{cd}} n_z^c \left( \tilde{p}_z^i \frac{\lambda_s^i}{\lambda_s^c} \right)^{1-\sigma_s} \right]^{\frac{1}{1-\sigma_s}} u_s^*. \quad (13)$$

We can now obtain an explicit expression for  $M_{s(cd)}^c$  in equation (11):

$$M_{s(cd)}^c = \frac{\left[ \sum_{z \in I_s^{cd}} n_z^c (\tilde{p}_z^c)^{1-\sigma_s} \right]^{\frac{1}{1-\sigma_s}}}{\left[ \sum_{z \in I_s^{cd}} n_z^c \left( \tilde{p}_z^d \frac{\lambda_s^d}{\lambda_s^c} \right)^{1-\sigma_s} \right]^{\frac{1}{1-\sigma_s}}} = \tilde{P}_s^{cd} \lambda_s^c \frac{\left[ \sum_{z \in I_s^{cd}} n_z^c \left( \frac{\tilde{p}_z^c}{\tilde{P}_s^c} \right)^{1-\sigma_s} \right]^{\frac{1}{1-\sigma_s}}}{\left[ \sum_{z \in I_s^{cd}} n_z^c \left( \frac{\tilde{p}_z^d}{\tilde{P}_s^d} \right)^{1-\sigma_s} \right]^{\frac{1}{1-\sigma_s}}} \quad (14)$$

Taking logarithms on both sides of (14) and using the fact that  $P_s^{cd} = \tilde{P}_s^{cd} \lambda_s^c$ , we can combine this equation with (11) to obtain

$$\ln H_s^{cd} \leq \ln M_{s(cd)}^c = \ln P_s^{cd} + \ln \phi_{s(cd)}^c, \quad \phi_{s(cd)}^c \equiv \frac{\left[ \sum_{z \in I_s^{cd}} n_z^c \left( \frac{\tilde{p}_z^c}{\tilde{P}_s^c} \right)^{1-\sigma_s} \right]^{\frac{1}{1-\sigma_s}}}{\left[ \sum_{z \in I_s^{cd}} n_z^c \left( \frac{\tilde{p}_z^d}{\tilde{P}_s^d} \right)^{1-\sigma_s} \right]^{\frac{1}{1-\sigma_s}}}. \quad (15)$$

Similarly, an expression analogous to (14) can be obtained for  $M_{s(cd)}^d$ , which combined with (12) yields<sup>16</sup>

$$\ln L_s^{cd} \geq \ln M_{s(cd)}^d = \ln P_s^{cd} + \ln \phi_{s(cd)}^d, \quad \phi_{s(cd)}^d \equiv \frac{\left[ \sum_{z \in I_s^{cd}} n_z^d \left( \frac{\tilde{p}_z^c}{\tilde{P}_s^c} \right)^{1-\sigma_s} \right]^{\frac{1}{1-\sigma_s}}}{\left[ \sum_{z \in I_s^{cd}} n_z^d \left( \frac{\tilde{p}_z^d}{\tilde{P}_s^d} \right)^{1-\sigma_s} \right]^{\frac{1}{1-\sigma_s}}}. \quad (16)$$

Equations (15) and (16) relate the implications of consumer cost minimization to cross-sectional Paasche and Laspeyres price indexes, where each of the cost-of-utility indexes has observable bounds on just one side. Our consideration of two cost-of-utility indexes, as well as the one-sidedness of their bounds, differs from the standard bounding

<sup>15</sup>It is here where Assumptions 1 and 2 are critical. In equation (13) we use these assumptions to derive  $\frac{p_z^i}{\lambda_z^c \xi_z^c} = \frac{p_z^i}{\lambda_z^i \xi_z^i} \frac{\lambda_z^i \xi_z^i}{\lambda_z^c \xi_z^c} = \tilde{p}_z^i \frac{\lambda_z^i}{\lambda_z^c}$ ,  $i = c, d$ .

<sup>16</sup>Note that all prices (observed and pure) in this section are *cif* import prices, that is, import prices inclusive of customs, insurance and freight charges. Under the assumption that trade costs are constant across product categories within a sector (see Section 4), the indexes  $M_{s(cd)}^c$ ,  $M_{s(cd)}^d$ ,  $H_s^{cd}$ ,  $L_s^{cd}$  alternatively can be defined as free-on-board (*fob*) – i.e., exclusive of customs, insurance and freight charges – if they are appropriately scaled by relative trade costs between countries  $c$  and  $d$  and the United States. As a result, the inequalities in equations (15) and (16) also apply to *fob* import prices. As noted in Section 5, we use *fob* import unit values to measure U.S. trading partners' export prices in our empirical analysis.

of cost-of-utility indexes from both above *and* below found in the index number literature. Here, since we allow for horizontal differentiation, we must deal with two cost-of-utility indexes because  $M_{s(cd)}^c$  and  $M_{s(cd)}^d$  are defined over different numbers of varieties, i.e.,  $n_z^c$  and  $n_z^d$ , respectively.<sup>17</sup> As a result,  $\phi_{s(cd)}^c$  and  $\phi_{s(cd)}^d$  are also different. Under plausible assumptions described below, however, we can show that  $\ln \phi_{s(cd)}^c < 0$  and  $\ln \phi_{s(cd)}^d > 0$ , which implies that the Paasche and Laspeyres indexes bound the Impure Price Index, i.e.,  $\ln H_s^{cd} \leq \ln M_{s(cd)}^c \leq \ln P_s^{cd} \leq \ln M_{s(cd)}^d \leq \ln L_s^{cd}$ .

### 3.2. Paasche and Laspeyres Bounds on the Impure Price Index

In this section we develop additional notation specific to country pairs. For each pair of countries  $c$  and  $d$ , define the pair's ( $o$ -normalized) average number of varieties in product category  $z$ :

$$\widehat{n}_z^{cd} = \frac{1}{2} \left( \frac{\bar{n}_s^o}{\bar{n}_s^c} n_z^c + \frac{\bar{n}_s^o}{\bar{n}_s^d} n_z^d \right), \quad (17)$$

and the country pair's ( $o$ -normalized) “multilateral excess variety” in product  $z$  relative to the world average:

$$\widetilde{n}_z^{cd} = \widehat{n}_z^{cd} - \bar{n}_z. \quad (18)$$

Multilateral excess variety measures the extent to which the average number of varieties in countries  $c$  and  $d$  is above or below the world average.

Also, for each country  $i = c, d$  in the country pair, define  $i$ 's ( $o$ -normalized) “bilateral excess variety” in product  $z$  relative to the country-pair average,

$$\widetilde{n}_z^{i,cd} = \frac{\bar{n}_s^o}{\bar{n}_s^i} n_z^i - \widehat{n}_z^{cd}. \quad (19)$$

Bilateral excess variety measures the extent to which the number of varieties in a country is above or below the bilateral average. These measures of excess variety possess three convenient properties:

$$\sum_z \widetilde{n}_z^{i,cd} = 0, \quad \sum_z \widetilde{n}_z^{cd} = 0, \quad \widetilde{n}_z^{c,cd} = -\widetilde{n}_z^{d,cd} \quad (20)$$

The first and second properties indicate that, across product categories within country  $i$ , both bilateral and multilateral excess variety sum to zero. The third property reveals that two countries cannot both have positive bilateral excess variety in the same category.

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<sup>17</sup>  $M_{s(cd)}^c$  and  $M_{s(cd)}^d$  would be equal, for example, if the number of varieties in countries  $c$  and  $d$  were proportional to one another for every product category.

Define the bilateral difference in two countries' pure prices in product category  $z$  relative to their countries' pure price aggregator as

$$\Delta \tilde{p}_z^{cd} = \left( \frac{\tilde{p}_z^c}{\tilde{P}_s^c} \right)^{1-\sigma_s} - \left( \frac{\tilde{p}_z^d}{\tilde{P}_s^d} \right)^{1-\sigma_s}. \quad (21)$$

A positive  $\Delta \tilde{p}_z^{cd}$  indicates that country  $c$  has a lower pure price of  $z$  (relative to the price aggregator) than country  $d$ . A lower pure price may arise, for example, due to comparative advantage, i.e., variation in exporters' relative production efficiency or factor costs.

Finally, for set of products  $A$ , define the sample covariance over that set of products as  $cov_A(x_z, y_z) = (1/Z_A) \sum_{z \in A} (x_z - \bar{x})(y_z - \bar{y})$ , where  $Z_A$  is the number of elements in  $A$ .

We now lay out a set of sufficient conditions for the Impure Price Index to be bounded by observable Paasche and Laspeyres price indexes.

Assumption 3 states that country  $c$  relative to country  $d$  will tend to have positive bilateral excess variety in those products in which it has a lower relative pure price.

**Assumption 3:**  $cov_{I_s^{cd}}(\tilde{n}_z^{c,cd}, \Delta \tilde{p}_z^{cd}) = cov_{I_s^{cd}}(\tilde{n}_z^{d,cd}, \Delta \tilde{p}_z^{dc}) \geq 0$

This assumption is based on the results of theoretical models of international trade with product differentiation that allow for trade costs and do not assume factor price equalization (e.g., Romalis 2004, Bernard et al. 2007). These models find that, across goods, the relative number of varieties between two countries is a negative function of the countries' relative prices. This finding supports the intuitive notion that countries should have a relatively higher (lower) number of firms in sectors or product categories in which they are relatively more (less) competitive, i.e. those sectors with relatively lower (higher) prices. It is possible to reformulate these models in terms of quality-adjusted variables. Thus reinterpreted, these models predict that the relative number of varieties in a sector or product category is a negative function of relative pure (or quality-adjusted) prices.

Assumption 4 imposes the restriction that there is no correlation between the country-pair's multilateral excess variety and bilateral differences in pure relative prices.

**Assumption 4:**  $cov_{I_s}(\tilde{n}_z^{cd}, \Delta \tilde{p}_z^{cd}) = 0$

This assumption is not very strong, as there is no obvious relationship between the country pair's excess variety relative to the world average and relative comparative advantage among countries *within* the pair.

Assumption 5 requires that countries  $c$  and  $d$  be similarly active in exporting goods to the United States.

**Assumption 5:**  $\delta_s^{cd} = \delta_s^{dc} = 0$ , where

$$\delta_s^{cd} = \frac{\sum_{z \in U_s^{cd}} \left( \frac{\sum_{z \in I_s^{cd}} \Delta \tilde{p}_z^{cd}}{Z_s^{cd}} + \hat{n}_z^{cd} \Delta \tilde{p}_z^{cd} \right)}{\sum_{z \in I_s^{cd}} n_z^c \left( \frac{\tilde{p}_z^d}{\tilde{p}_s^d} \right)^{1-\sigma_s}}, \text{ and}$$

$\delta_s^{dc}$  is defined analogously.

The magnitude of the terms  $\delta_s^{cd}$  and  $\delta_s^{dc}$  depends on the extent to which countries  $c$  and  $d$  are “similarly active”. Assumption 5 requires that these terms are zero. A sufficient condition that implies assumption 5 is that the two countries are active in the same categories. In that case, the numerator in the expression for  $\delta_s^{cd}$  is zero, as it sums over elements of an empty set,  $U_s^{cd}$ . Since the sums in the numerator involve positive and negative terms, it is still possible that the numerator is zero even if  $U_s^{cd}$  is non-empty. More generally,  $\delta_s^{cd}$  and  $\delta_s^{dc}$  will tend to be smaller (in absolute magnitude) the smaller is the number of mismatched active categories (in the numerator) relative to the number of matched active categories (in the denominator).<sup>18</sup>

With assumptions 3, 4 and 5 as well as our earlier assumptions about consumer utility, Proposition 1 demonstrates that a country pair’s unobservable Impure Price Index is bounded by the observable Paasche and Laspeyres indexes defined over their common exports to a third country.

**Proposition 1** *Under Assumptions 1 through 5, for any two countries  $c$  and  $d$ , the (unobservable) Impure Price Index is bounded by the (observable) Paasche and Laspeyres indexes:*

$$\ln H_s^{cd} \leq \ln P_s^{cd} \leq \ln L_s^{cd}$$

**Proof.** See Theory Appendix. ■

This finding provides the basis for our estimation of the Impure Price Index in the first-stage of our empirical strategy.

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<sup>18</sup>The empirical section imposes a threshold number of matched active categories for a country pair to be included in the estimation.

#### 4. Net Trade as Indicator of Pure Price Variation

This section derives the theoretical relationship between countries' net trade and their Pure Price Indexes. Exporting goods from country  $c$  to country  $c'$  requires paying iceberg trade costs of  $\tau_s^{cc'}$ . Therefore,  $p_z^c \tau_s^{cc'}$  is the import price of product  $z$  in country  $c'$ . Given the CES preference structure assumed in equation (1), it is easy to derive country  $c$ 's bilateral export and import flows (in sector  $s$ ) with every other country  $c'$ . Summing export flows over  $c' \neq c$ , we obtain the value of country  $c$ 's exports,

$$Exports_s^c = \sum_{c' \neq c} \left[ \sum_z \frac{n_z^c (\tilde{p}_z^c \tau_s^{cc'})^{1-\sigma_s}}{(G_s^c)^{1-\sigma_s}} \right] b_s Y^c \quad (22)$$

where  $Y^c$  is the income of country  $c$  and

$$G_s^c = \left[ \sum_{c''} \sum_z n_z^{c''} (\tilde{p}_z^{c''} \tau_s^{c''c'})^{1-\sigma_s} \right]^{1/(1-\sigma_s)} \quad (23)$$

is a consumption-based price aggregator. The expression in brackets in equation (22) is country  $c$ 's share in country  $c'$ 's sectoral expenditure,  $b_s Y^c$ . This share does not depend on prices and quality levels independently of one another, but only on the ratio of the two,  $\tilde{p}_z^c$ .<sup>19</sup>

In a similar manner, we obtain the value of country  $c$ 's imports,

$$Imports_s^c = \sum_{c' \neq c} \sum_z \frac{n_z^{c'} (\tilde{p}_z^{c'} \tau_s^{c'c})^{1-\sigma_s}}{(G_s^c)^{1-\sigma_s}} b_s Y^c = \left[ 1 - \sum_z \frac{n_z^c (\tilde{p}_z^c)^{1-\sigma_s}}{(G_s^c)^{1-\sigma_s}} \right] b_s Y^c. \quad (24)$$

Subtracting equation (24) from equation (22), we obtain country  $c$ 's net trade with the world in sector  $s$ ,  $T_s^c$ , as a proportion of its expenditure in the sector,

$$\frac{1}{b_s} \frac{T_s^c}{Y^c} = -1 + \sum_{c'} \sum_z \frac{n_z^c (\tilde{p}_z^c \tau_s^{cc'})^{1-\sigma_s}}{(G_s^{c'})^{1-\sigma_s}} \frac{Y^{c'}}{Y^c}, \quad (25)$$

Equation (25) shows that countries' trade balance in sector  $s$  is a function of all the product-level pure prices in that sector. Our objective is to simplify this expression by relating net trade of country  $c$  in sector  $s$  to the Pure Price Index.

To express equation (25) as a function of the Pure Price Index, we must impose structure on the relationship between pure prices and the number of varieties countries

<sup>19</sup>We can associate an infinite price  $\tilde{p}_z^c$  with a product  $z$  that is not produced in country  $c$ . Since pure prices are elevated to a negative exponent, this product will have no effect on the volume of trade or the price aggregator.

produce. Based on the same theoretical results that motivate Assumption 3, we postulate a negative relationship between the number of varieties and pure prices, defined here across sectors rather than across categories within sectors.

**Assumption 6:**  $\frac{\bar{n}_s^c/Y^c}{\bar{n}_s^o/Y^o} = \left(\tilde{P}_s^{co}\right)^{-\eta_s}$ ,  $\forall c = 1, \dots, C$ ,  $\eta_s \geq 0$ .

A particular case of this assumption is when  $\eta_s = 0$ , in which case the average number of varieties in a sector is a constant proportion of income. Here, we allow for a more general case where the number of varieties is allowed to decrease as pure prices increase. We also characterize the relationship between pure prices and number of varieties across product categories within sectors as the sum of a common component across countries ( $V_s$ ) and a mean-zero, country-specific idiosyncratic component

$$\text{cov} \left[ \tilde{n}_z^c, \left( \tilde{p}_z^c / \tilde{P}_s^c \right)^{1-\sigma_s} \right] = V_s + \mu_s^c, \quad (26)$$

noting that this characterization does not impose any restriction on the covariance.<sup>20</sup>

The following Proposition describes the main result of this section.

**Proposition 2** *Under Assumption 6, country  $c$ 's sectoral net trade can be expressed (abstracting from approximation error) as a linear function of the Pure Price Index*

$$\frac{T_s^c}{Y^c} = \Psi_s + \gamma_s \ln \tilde{P}_s^{co} + b_s \tau_s^c + \iota_s^c \quad (27)$$

where

$$\begin{aligned} \gamma_s &= b_s(1 - \sigma_s - \eta_s) < 0, & \Psi_s &= b_s(k_s + Z_s V_s), \\ \tau_s^c &= \ln \left( \sum_{c'} Y^{c'} \left( \frac{\tau_s^{cc'}}{G_s^{c'}} \right)^{1-\sigma_s} \right), & k_s &= \ln \left[ \frac{(\tilde{P}_s^o)^{1-\sigma_s}}{Y^o} \right], \\ \iota_s^c &= b_s Z_s \mu_s^c \end{aligned}$$

**Proof.** See Theory Appendix. ■

Proposition 2 provides a simple expression for the relationship between net trade and pure prices. This proposition formalizes the key insight of the paper. Price variation not accompanied with corresponding quality variation implies variation in pure prices. Even though unobservable, pure prices are manifest in sectoral trade balances. In particular, the surplus in a country's sectoral net trade should be larger the lower are its pure prices.

In addition to pure prices, trade costs also influence net trade. The term  $\tau_s^c$  summarizes this influence. This term includes bilateral trade costs between all country pairs. First, it

<sup>20</sup>For estimation, we will assume that  $\mu_s^c$  and the instrumental variable are uncorrelated.

includes all outbound bilateral trade costs for country  $c$ . Those costs,  $\tau_s^{cc'}$ , enter directly into the summary term  $\tau_s^c$ , which is smaller the higher are those costs. Second,  $\tau_s^c$  also includes all inbound bilateral trade costs for country  $c$ ,  $\tau_s^{c'c}$ . In this case, the inbound costs enter through the consumption price index for country  $c$ ,  $G_s^c$ . The term  $\tau_s^c$  is larger the larger are those costs, and hence affect positively net trade of country  $c$ . Finally, all other bilateral trade costs enter indirectly through countries' consumption price indexes,  $G_s^{c'}$ , dampening the negative effect of the outbound bilateral trade costs. Therefore, net trade of country  $c$  is higher the higher are trade costs between third countries.<sup>21</sup>

The impact of trade costs on net trade characterized in Proposition 2 is *conditional on pure prices*. That is, while trade costs properly shift the relationship between net trade and pure prices, they do not provide a comparative statics assessment of the effect of trade costs on net trade. Changes in those costs will typically affect pure prices in general equilibrium, implying an indirect effect on net trade not captured in equation (27). Note that our method does not require that we identify the economic forces that determine pure prices in equilibrium. It only requires that we control for them. Variation in pure prices can be driven by traditional sources of comparative advantage, or it can be the result of macroeconomic conditions, such as over- or under-valued currencies.

Equation (27) can be interpreted as a relative demand function, where net trade is the “quantity” variable, the Pure Price Index is the “price” variable, and the trade costs are demand shifters. The first term captures movements along the demand curve: higher pure prices of country  $c$  in sector  $s$  are associated with a worsening of this country's net trade position in that sector. The second term captures movements of the demand curve. Conditional on pure prices, higher inbound trade costs relative to outbound trade costs shift this curve to the right.

Before concluding this section, we note that a substantial advantage of using net trade as the quantity indicator in our method is that it can mitigate our inability to control for components of trade costs we cannot observe, such as information costs and non-tariff barriers associated with commercial policy. As noted above, bilateral trade costs affect countries' exports and imports in opposite directions. As a result, the impact of unobserved trade costs may, to a considerable extent, cancel out. In contrast, an alternative approach based purely on “demand” information, e.g., using countries' U.S. import market shares rather than their global net trade, as in Khandelwal (2007), is likely to be substantially more sensitive to mismeasurement of trade costs.

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<sup>21</sup>See Anderson and van Wincoop (2003) for a detailed discussion of the effects of trade costs on trade flows in a related setting.

## 5. Estimation

In this section we demonstrate how our theoretical results can be used to estimate U.S. trading partners' relative manufacturing export quality from 1989 to 2003. Estimation is accomplished in two stages. We discuss the strategy of each stage, as well as their data requirements, separately. Throughout, we focus on the key issues associated with implementing our method, deferring detailed discussions of data requirements to a separate Data Appendix. Raw datasets and the computer code used to generate the results reported below are available from the authors upon request.

### 5.1. Estimation of First-Stage Impure Price Indexes

The first stage of the estimation uses Proposition 1 to estimate each country's Impure Price Index,  $\hat{P}_s^{ko}$ ,  $\forall k \neq o$ , where country  $o$  is the numeraire country and hats over variables denote estimates. We note that the choice of numeraire is made without loss of generality; in our empirical implementation, we use Switzerland. For generic country pair  $c$  and  $d$ , the estimated indexes  $\hat{P}_s^{co}$  and  $\hat{P}_s^{do}$  implicitly determine the bilateral index  $\hat{P}_s^{cd} = \hat{P}_s^{co} / \hat{P}_s^{do}$ . This index should satisfy the Paasche and Laspeyres bounds for that country pair, as outlined in Proposition 1. Similarly, for  $C$  trading partners, the  $C - 1$  estimated Impure Price Indexes  $\hat{P}_s^{ko}$ ,  $\forall k \neq o$ , implicitly determine  $C(C - 1)$  bilateral indexes  $\hat{P}_s^{cd}$ ,  $\forall c, d$ , that should satisfy the bilateral Paasche and Laspeyres bounds for all country pairs.

If export prices and quantities were observed without error, estimation would entail searching for an interior solution to the set of observed Paasche and Laspeyres bounds across country pairs. Given that import data may be mis-recorded on customs documents, however, we allow for measurement error in the bounds by assuming that Paasche and Laspeyres indexes are observed imprecisely. Denote the "true" Paasche and Laspeyres indexes by  $H_s^{*cd}$  and  $L_s^{*cd}$ , respectively. We assume that the observed indexes,  $H_s^{cd}$  and  $L_s^{cd}$ , depart from the true indexes by a multiplicative error: in logs,  $\ln H_s^{cd} = \ln H_s^{*cd} + \varrho_{h,s}^{cd}$  and  $\ln L_s^{cd} = \ln L_s^{*cd} + \varrho_{l,s}^{cd}$ . We also assume that each error is distributed normally, with mean zero and standard deviation  $\psi_s$ , and that the errors for each bound are independent both of each other and of error terms for other bilateral pairs.<sup>22</sup>

Satisfying the inequality constraints of Proposition 1 for a given pair of countries

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<sup>22</sup>This is a potentially strong assumption because the price (unit value) of a single product might show up in many bounds, inducing correlated rather than independent errors. It is worth noting that biases in the estimated standard errors of the estimated price indexes do not affect our results for quality since only price index point estimates are used in the second stage. Biases in the latter estimates, however, may not be innocuous.

implies:

$$\ln P_s^{cd} \geq \ln H_s^{*cd} \Rightarrow \varrho_{h,s}^{cd} \geq \ln H_s^{cd} - \ln P_s^{cd} \quad (28)$$

$$\ln P_s^{cd} \leq \ln L_s^{*cd} \Rightarrow \varrho_{l,s}^{cd} \leq \ln L_s^{cd} - \ln P_s^{cd}. \quad (29)$$

Separately for each year  $t$ , we estimate a set of index numbers  $\ln \widehat{P}_s^{ko}$ ,  $\forall k \neq o$ , and the standard deviation of the error term  $\widehat{\psi}_s$  by maximizing the joint likelihood that the intervals defined by all “true” Paasche and Laspeyres bounds contain the estimates, i.e. the likelihood that (28) and (29) are jointly satisfied for each country pair  $\{c, d\}$ . This criterion implies maximizing the function

$$\ln L = \sum_c \sum_{d>c} \left\{ \ln \left[ 1 - \Phi \left( \frac{\ln H_s^{cd} - \ln P_s^{cd}}{\psi_s} \right) \right] + \ln \Phi \left( \frac{\ln L_s^{cd} - \ln P_s^{cd}}{\psi_s} \right) \right\}$$

where  $\Phi$  is the cumulative normal.

Intuition for this estimator is provided in Figure 1, which considers the Paasche-Laspeyres interval for a single country pair  $c$  and  $d$ , defined by  $\ln H_s^{cd}$  and  $\ln L_s^{cd}$ . In the figure, two cumulative normal distributions, each with standard deviation  $\psi_s$ , take values of one half at each end of the interval. Consider a pair of Impure Price Index estimates relative to the numeraire and the location of their (log) ratio  $\ln \widehat{P}_s^{cd} = \ln \widehat{P}_s^{co} - \ln \widehat{P}_s^{do}$  along the horizontal axis in the figure. According to equation (28), the height of the cumulative normal distribution to the left of  $\ln \widehat{P}_s^{cd}$  indicates the likelihood that the true Paasche index is lower than the estimated bilateral index, that is,  $\ln H_s^{*cd} < \ln \widehat{P}_s^{cd}$ . Likewise, using equation (29), the height of the cumulative normal to the right of  $\ln \widehat{P}_s^{cd}$  indicates the likelihood that the true Laspeyres index is greater than the estimated bilateral index, that is,  $\ln L_s^{*cd} > \ln \widehat{P}_s^{cd}$ . Choosing a particular value for  $\ln \widehat{P}_s^{cd}$  inevitably involves increasing the value of one of these functions at the expense of the other. If the objective were to maximize the likelihood that  $\ln \widehat{P}_s^{cd}$  is within the true bilateral Paasche and Laspeyres bounds, only taking into account the bounds of this particular country pair, then  $\ln \widehat{P}_s^{cd}$  would lie in the middle of the interval and be equivalent to the well-known Fisher index. However, because the choices of  $\ln \widehat{P}_s^{co}$  and  $\ln \widehat{P}_s^{do}$ , which determine  $\ln \widehat{P}_s^{cd}$  for this country pair, also influence the fit of all other country pairs in which either country  $c$  or  $d$  are present, the estimates that maximize the joint likelihood for all country pairs will not in general be located in the center of the interval for countries  $c$  and  $d$ . For that reason,  $\ln \widehat{P}_s^{cd}$  is drawn off-center in the interval depicted in Figure 1. This estimator has the advantage that it penalizes estimates that lie inside the interval only in relation to the likelihood that conformance to the theory is a consequence of measurement error. Similarly, it

penalizes estimates outside the interval only in relation to the likelihood that violation of the bounds restriction is not caused by measurement error. We note that this estimator is not a conventional maximum likelihood estimator as it does not maximize the likelihood of observing the data (the bounds) given the parameters (the Impure Price Indexes).

Even though the theoretical properties of our estimator are unknown, further intuition for it can be derived from consideration of two alternatives. The first, which we refer to as the “V” estimator, is defined by a quadratic penalty function centered at the midpoint ( $F_s^{cd}$ ) of each country pair’s interval,

$$\Psi^V = \sum_c \sum_{d>c} (\ln F_s^{cd} - \ln P_s^{cd})^2. \quad (30)$$

Since the midpoint of the interval is equal to the (log of the) Fisher index defined by the intervals’ Paasche and Laspeyres bounds ( $F_s^{cd} = \sqrt{H_s^{cd}L_s^{cd}}$ ), this penalty function is similar in spirit to other multilateral indexes proposed in the index number literature (see, for example, Diewert and Nakamura 1993). Though this approach has the advantage of rewarding estimates that are closer to the middle of the interval, where conformance with the bounds is less likely to be driven by measurement error, it has the undesirable feature of treating equally deviations from the Fisher index that are inside versus outside of the theoretically mandated bounds.

A second alternative penalty function, which we refer to as the “sink” estimator, only penalizes estimates outside the interval:

$$\Psi^S = \sum_c \sum_{d>c} 1_s^{cd} * [\min \{abs(\ln P_s^{cd} - \ln H_s^{cd}), abs(\ln L_s^{cd} - \ln P_s^{cd})\}]^2 \quad (31)$$

where  $1_s^{cd}$  is an indicator variable equalling zero for  $\ln H_s^{cd} \leq \ln P_s^{cd} \leq \ln L_s^{cd}$  and one otherwise. While this approach properly favors estimates within the interval, it ignores potential measurement error. Our proposed estimator, by contrast, penalizes estimates within and outside the interval, but only according to the likelihood that conformance to the theory is a consequence of measurement error.

We also investigate the use of an index (HK) proposed by Hummels and Klenow (2005),

$$HK_s^{kW} = \prod_{z \in I_s^{kW}} \left( \frac{p_z^k}{p_z^W} \right)^{w_z^k}, \quad k = 1, \dots, C \quad (32)$$

where

$$s_z^k = \frac{p_z^k q_z^k}{\sum_{z' \in I_s^{kW}} p_{z'}^k q_{z'}^k}, \quad s_z^W = \frac{p_z^W q_z^W}{\sum_{z' \in I_s^{kW}} p_{z'}^W q_{z'}^W}, \quad w_z^k = \frac{\frac{s_z^k - s_z^W}{\ln s_z^k - \ln s_z^W}}{\sum_{z' \in I_s^{kW}} \frac{s_{z'}^k - s_{z'}^W}{\ln s_{z'}^k - \ln s_{z'}^W}}. \quad (33)$$

The index  $HK_s^{kW}$  compares country  $k$ 's prices to those of the world ( $W$ ) over the set of goods country  $k$  has in common with the world ( $z \in I_s^{kW}$ ). The world ( $W$ ) is an aggregate of all countries – in our case the 43 countries in the sample – except for country  $k$ . Therefore, the “world” price of product  $z$ ,  $p_z^W$ , is just total world value divided by total world quantity,  $q_z^W$ , omitting country  $k$  in the calculation.<sup>23</sup> Though this index has the advantage that it can be computed rather than estimated, it has the disadvantage of treating an aggregation of countries as a single entity without theoretical justification. Finally, a bilateral index for country pair  $c$  and  $d$ ,  $HK_s^{cd}$ , is computed as the ratio  $HK_s^{cW}/HK_s^{dW}$ .<sup>24</sup>

We compare the empirical performance of each of these estimators after outlining the first-stage data requirements. As discussed further below, we show that our preferred estimator comes closest to estimating Impure Price Indexes consistent with Proposition 1.

#### 5.1.1. First-Stage Data Requirements

Estimation of countries' Impure Price Indexes requires data on countries' export prices and quantities. Here, we rely on detailed U.S. import statistics published by the U.S. Census Bureau. These data record the total customs value and quantity of U.S. imports by year, source country and ten-digit Harmonized System (HS) product classification from 1989 to 2003. We focus here on U.S. import data given its availability for such a long time horizon, but note that our method can be generalized to include data from other countries; such data would generate additional Paasche and Laspeyres bounds that could be incorporated into the estimation. Our use of U.S. trade data presumes that U.S. import prices and quantities are representative of countries' exports to other markets.<sup>25</sup>

We compute the unit value, or “price”, of export product  $z$  from source country  $c$ ,  $p_z^c$ , by dividing free-on-board import value ( $v_z^c$ ) by import quantity ( $q_z^c$ ),  $p_z^c = v_z^c/q_z^c$ , where free-on-board refers to import values that are exclusive of customs, insurance and freight charges.<sup>26</sup> Examples of the units employed to classify products include dozens of men's

<sup>23</sup>Since country  $k$  is not included to calculate world prices, the set of countries in  $W$  varies with  $k$ . We do not subindex  $W$  by  $k$  to simplify notation.

<sup>24</sup>Note that since the  $HK$  index is a bilateral index applied to a multilateral purpose, it does not satisfy transitivity, i.e., it cannot be obtained from applying equation (32) directly to countries  $c$  and  $d$ . Formally,  $HK_s^{cd} \neq \prod_{z \in I_s^{cd}} \left( \frac{p_z^c}{p_z^d} \right)^{w_z^{cd}}$ .

<sup>25</sup>In principle, this assumption could be tested by comparing the results of this section to results based on other countries' data.

<sup>26</sup>A sustained assumption in our framework is that the export unit values that we observe are not systematically different from the prices charged to domestic consumers, which we do not observe.

cotton shirts in apparel, square meters of wool carpeting in textiles and pounds of folic acid in chemicals. We focus on manufacturing exports, where a product is classified as manufacturing if it belongs to SITC industries 5 through 8. Following standard practice, we exclude SITC 68, non-ferrous metals, from manufacturing. We note that quantity information is missing for approximately 20 percent of observations in the raw data; these observations are dropped.

Unit values are noisy due to both aggregation and measurement errors (GAO 1995). To mitigate the impact of these errors, we both restrict our analysis to relatively large exporters and screen the raw data. First, we start with the world’s top 50 exporters of manufactured goods by value. Second, we employ two types of screens to eliminate suspect observations. “Primary” screening drops observations where only a single unit is shipped in a year or where the U.S. CPI-deflated annual import value is below \$25,000 in 1989 dollars. “Secondary” screening makes the primary quantity and value cutoffs more stringent while imposing three additional criteria:

- *(More Stringent) Relevance Constraint:* Country-product-year observations must have quantity greater than 25 and value (in 1989 dollars) greater than \$50,000;
- *Presence Constraint:* Country-product observations must appear in more than two years of the sample;
- *Country-Pair Overlap Constraint:* For a country-pair comparison to be included in the sample in any given year, the two countries must export at least 25 products in common to the United States; and
- *Unit-Value Dispersion Constraint:* Country-product-year observations are excluded if the country’s adjusted<sup>27</sup> unit value is less than one-fifth or more than five times the geometric mean of all prices for the product in that year.

After secondary-screening the data, we finally impose the constraint that data required for both the first and second stage cannot be missing for more than three years of the

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<sup>27</sup>The adjustment accounts for the likelihood that very high export prices are more likely to be the result of misrecording if they come from countries with relatively low average export prices, and *vice versa*. To adjust product-country-year unit values relative to the numeraire country, we perform two iterations of the first-stage estimation. In the first iteration, we estimate Impure Price Indexes after eliminating observations under the unit-value-dispersion constraint without making any adjustment to country’s unit values. In the second iteration, we divide a country’s unit values by the estimated Impure Price Index from the first iteration prior to implementing the unit-value-dispersion screen. We note that omitting the second iteration has relatively little impact on our second-stage quality estimates.

sample period. After all screens are implemented, we are left with 43 countries, which constitute the base sample we use in the remainder of the paper.

The costs and benefits of screening the raw data can be discerned from Table 1. Each row of the table focuses on a different screen, while each column indicates the affect of the screen on a different aspect of the 2003 sample, though we note that screening has a similar effect across years. To promote comparability, all rows in the table are restricted to the same set of 43 countries available after the most stringent screening (that is, the screening in the final row of the table).

The first column of Table 1 demonstrates that secondary screening reduces the value of imports captured in the sample by 11 percent *vis a vis* the primary-screened sample. The next two columns of Table 1 show that secondary screening also reduces country and country-product participation in the sample, lowering the number of country pairs for which data is available to 829 from 861 and the median number of products country pairs export in common to the United States from 347 to 228. As illustrated in the final column of the table, there are very few incorrectly ordered Paasche and Laspeyres bounds (i.e.,  $L_s^{cd} < H_s^{cd}$ ) in all three screens. We do not use those bounds for estimation.

The primary benefit of screening is substantially tighter Paasche and Laspeyres bounds. As indicated in the fourth column of the table, the median interval length ( $\ln L_s^{cd} - \ln H_s^{cd}$ ) under the preferred secondary screening is 0.74, less than one-third the length under the primary screen, 2.51. The reduction in interval length results in a substantial improvement in estimation precision.

Of the additional criteria imposed by secondary screening, the unit-value dispersion constraint exerts the strongest affect on median interval length. For example, an “alternate” secondary screening (not shown) that omits the requirement that adjusted unit values be within one-fifth and five times the geometric mean for the product-year results in a disproportionately large increase in median interval length (to 2.01) versus import value (to 97.8 percent).

The left-hand panel of Table 2 summarizes several dimensions of the preferred sample, by year. The first column of the panel illustrates that the sample of countries is held constant at 43 for the entire sample period. The final column of the panel shows that the median Paasche-Laspeyres interval across country pairs measured in log points moves between 0.68 and 0.78 over the sample period. The remaining columns of the panel demonstrate that the number of country pairs, the total number of product-country-pairs, and the median number of common products across country pairs all rise over time. These increases are driven by growth in the number of products countries export to the

United States over the sample period. This growth introduces a potential composition bias into our estimates, a well-known problem in the index number literature. We attempt to mitigate the influence of composition bias via use of the “presence” and “overlap” screens outlined above. Unfortunately, we cannot restrict our analysis to a set of continually exported country-products due to numerous changes in product classification codes during the sample period.

### 5.1.2. First-Stage Results

The right-hand panel of Table 2 summarizes the results of the first-stage estimation by year. Column one of the panel shows that the log likelihood declines in absolute value over time, while column two reports that the estimated standard deviation,  $\hat{\psi}_s$ , is relatively constant at approximately 0.15 over the sample period. The third column of the panel reports the estimation’s goodness of fit in terms of the percent of first-stage Impure Price Index estimates that lie within the Paasche-Laspeyres bounds. As indicated in the table, this share is above 90 percent in all years and rises from 90.4 percent in 1989 to 93.8 percent in 2003.

Goodness of fit for the alternate “V”, “sink” and HK Impure Price Index estimates described above, in contrast, is generally lower, supporting the choice of our preferred approach. As reported in Table 3, the “V” performs best among the alternatives to the preferred estimator, with the “sink” estimator a close second. The performance of the “HK” estimator, on the other hand, is poor: on average, just 43 percent of the bilateral Impure Price Indexes lie within the theoretically mandated Paasche and Laspeyres bounds. Similar differences are manifest in the first-stage estimates: though we find a high cross-country correlation between Impure Price Indexes estimated by our preferred estimator and those estimated by the “V” and the “sink” (the average cross-sectional correlation across years is above 0.99 in both cases), the correlation with the computed “HK” indexes is much lower (an average across years of 0.43).<sup>28</sup> Given the similarity of the preferred, “V” and “sink” estimates, it is not surprising that the second-stage quality estimates to which they give rise are also quite similar.

Estimation of the first stage yields an Impure Price Index for each country relative to the numeraire country. In Figure 2, we report *normalized* log Impure Price Indexes for all countries for the first and last years of the sample. This normalization involves

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<sup>28</sup>In a few cases, for a given country-year the “sink” estimator yields an indeterminate solution over a compact interval. This indeterminacy occurs for one country per year on average. Choosing alternative points within the interval has negligible effects on the cross-country correlations cited in the text.

subtracting the mean log index across countries from every country's estimated log Impure Price Index, by year

$$\ln \widehat{P}_{st}^{c,Mean} = \ln \widehat{P}_{st}^{co} - \frac{1}{C} \sum_k \ln \widehat{P}_{st}^{ko}. \quad (34)$$

In particular, the normalized Impure Price Index for the numeraire country,  $\ln \widehat{P}_{st}^{o,Mean}$ , is equal to  $-\frac{1}{C} \sum_k \ln \widehat{P}_{st}^{ko}$ .

Estimated Impure Price Indexes generally accord with expectations. In the figure, countries nearer the origin such as Pakistan (PAK) and China (CHN) exhibit relatively low export prices in both years *vis a vis* the mean while countries in the upper right corner like Ireland (IRL) and Switzerland (CHE) exhibit consistently high relative export prices. Countries' orientation with respect to the grey forty-five degree line illustrates changes in relative prices over time. Countries like Hungary and Morocco (MAR) that lie above the forty-five degree line exhibit rising relative export prices, while those below the forty-five degree line like China (CHN) and Singapore (SGP) experience declining relative prices. In both years, the ordering of countries accords well with their level of development. Note that a mapping of country codes to country names is provided in Table 5.

## 5.2. Estimation of Second-Stage Quality Indexes

The second stage of our estimation uses the result in Proposition 2 to recover information about countries' relative export quality from their first-stage estimated Impure Price Indexes. Given  $\ln \widetilde{P}_s^{cd} = \ln P_s^{cd} - \ln \lambda_s^{cd}$  from equation (7), we can rewrite the result of Proposition 2 as

$$T_{st}^c/Y_t^c = \Psi_{st} + \gamma_s \ln \widehat{P}_{st}^{co} + b_s \tau_{st}^c - \gamma_s \ln \lambda_{st}^{co} + \gamma_s \kappa_{st}^{co} + b_s Z_s \mu_{st}^c \quad (35)$$

where  $t$  indexes time periods and  $\kappa_{st}^{co} = \ln P_{st}^{co} - \ln \widehat{P}_{st}^{co}$  is the estimation error from the first stage. The last three terms in equation (35) are unobservable and create a compound error term that includes: countries' product quality relative to the numeraire country ( $\lambda_{st}^{co}$ ); the estimation error in the first stage ( $\kappa_{st}^{co}$ ); and the idiosyncratic component of the covariance between excess variety and pure prices ( $\mu_{st}^c$ ) from equation (26). Assuming that this compound error term is uncorrelated with the regressors is untenable. In particular, the quality component  $\lambda_{st}^{co}$  may be correlated with the estimated Impure Price Index: developed countries, which tend to have higher export prices, are also likely to produce higher quality.

To deal with this endogeneity, we first specify a linear time path for the evolution of product quality relative to the numeraire country,

$$\ln \lambda_{st}^{co} = \alpha_{0s}^{co} + \alpha_{1s}^{co}t + \varepsilon_{st}^{co} \quad (36)$$

where  $\alpha_{0s}^{co}$  and  $\alpha_{1s}^{co}$  are a country fixed effect and the slope of a country-specific time trend, respectively, and  $\varepsilon_{st}^{co}$  represents deviations of quality from this trend. As in the estimation of the first stage, results here do not depend upon the choice of numeraire country, and we again choose Switzerland for this role. Incorporating the country-specific linear time trend for quality into equation (35), we obtain our second-stage estimating equation

$$T_{st}^c/Y_t^c = \Psi_{st} + \gamma_s \ln \widehat{P}_{st}^{co} - \zeta_{0s}^{co} - \zeta_{1s}^{co}t + b_s \tau_{st}^c + v_{st}^{co} \quad (37)$$

where  $\zeta_{0s}^{co} = \gamma_s \alpha_{0s}^{co}$  and  $\zeta_{1s}^{co} = \gamma_s \alpha_{1s}^{co}$  are country fixed effects and time trends, respectively, and  $v_{st}^{co} = \gamma_s (\kappa_{st}^{co} - \varepsilon_{st}^{co}) + b_s Z_s \mu_{st}^c$  is the error term. Note that the terms for both the trade balance and the tariff could be expressed relative to the numeraire country, but that doing so would have an impact only on the year fixed effects.

The inclusion of country fixed effects in (37) eliminates the most obvious source of endogeneity, i.e. the cross-sectional correlation between the time-invariant components of countries' prices and quality levels. The inclusion of country-specific time trends further reduces the remaining correlation between regressor and error term, as the latter term now includes only deviations of quality from country-specific trends. However, correlation between  $\varepsilon_{st}^{co}$  and  $\widehat{P}_{st}^{co}$  may still persist, as shocks to quality may be accompanied by increases in (impure) prices.

To address this potential problem, we use the real exchange rate as an instrument for the estimated Impure Price Index. As usual, the instrument needs to satisfy two conditions. First, because the estimating equation includes country-specific fixed effects and time trends, the instrument has to be correlated with  $\ln \widehat{P}_{st}^{co}$ , *after controlling for the fixed effects and time trends*. In other words, deviations of the real exchange from its own time trend have to be correlated with similar deviations of  $\widehat{P}_{st}^{co}$ . Macroeconomic conditions typically determine periods of over- and under-valuation of countries' real exchange rate around long-run trends. These periods also determine changes in the international competitiveness of a countries' exports, captured in our model by  $\widetilde{P}_{st}^{co}$ . Since  $\widetilde{P}_{st}^{co}$  is a component of  $\widehat{P}_{st}^{co}$ , periods of over- or under-valuation are also associated with movements of  $\widehat{P}_{st}^{co}$ , providing the necessary correlation. Second, the instrument has to be uncorrelated with the error term  $\varepsilon_{st}^{co}$ , which requires that shocks to quality around the trend in sector  $s$  are not correlated with the real exchange rate. While we cannot rule out that such a

correlation exists, we judge it to be relatively unimportant. Shocks to quality in sector  $s$  might be accompanied by exactly offsetting changes in prices, leaving pure prices – and hence net trade in that sector – unchanged. Even if these shocks affect pure prices, they might have a negligible effect on the real exchange rate. This is more likely to be true if the shocks are temporary deviations around a trend, and if they are specific to sector  $s$ , that is, uncorrelated with shocks to quality in other sectors. Finally, we also assume that both  $\kappa_{st}^{co}$  and  $\mu_{st}^c$  are uncorrelated with the real exchange rate.

We estimate equation (37) using two-stage least squares (2SLS). Our estimate of country  $c$ 's Quality Index relative to the numeraire is

$$\ln \widehat{\lambda}_{st}^{co} = - \left( \frac{\widehat{\zeta}_{0s}^{co} + \widehat{\zeta}_{1s}^{co} t}{\widehat{\gamma}_s} \right), \quad (38)$$

where  $t$  indexes the number of years since 1989 and the remaining right-hand side variables are estimates from equation (37). Note that we identify only the linear trend in quality. Deviations of quality from the linear trend are confounded with the other two components of the error term and are therefore not included in our estimated Quality Indexes.

Countries' estimated Pure Price Indexes are derived from equation (37) and the definition of  $\ln \widehat{\lambda}_{st}^{co}$  in equation (38). They are equal to

$$\ln \widehat{P}_{st}^{co} = \ln \widehat{P}_{st}^{co} - \ln \widehat{\lambda}_{st}^{co} = \left( \frac{\frac{T_{st}^c}{Y_t^c} - \widehat{\Psi}_{st} + \widehat{b}_s T_{st}^c - \widehat{v}_{st}^{co}}{\widehat{\gamma}_s} \right). \quad (39)$$

We note that this estimate of the Pure Price Index inherits any estimation error in both the Impure Price Index and the Quality Index. In particular deviations of quality from the trend ( $\varepsilon_{st}^{co}$ ) are misattributed to the Pure Price Index.

### 5.2.1. Second-Stage Data Requirements

Estimation of the second stage faces a number of practical obstacles. Foremost among them is data collection. Obtaining reliable information about countries' trade balances, for example, is challenging because countries vary greatly in how they report this information to international agencies. Similarly, collection of countries' product-level trade barriers did not begin in earnest until 1989 and has grown fitfully since then. Here, we provide a brief description of how our datasets are constructed. See the separate Data Appendix for further detail.

Trade balance data are drawn from the United Nations Commodity Trade Statistics Database (COMTRADE). This dataset records bilateral import and export flows between

countries by manufacturing industry and year. Our overall approach to obtaining countries' net trade is to subtract each country's total reported imports from its total reported exports by industry and year.<sup>29</sup> We measure countries' annual net trade in overall manufacturing as well as the industries within manufacturing discussed below. As required by equation (37), we normalize trade balances by nominal GDP drawn from the World Bank's World Development Indicators (WDI) database. Data on Taiwan's GDP are from the Economist Intelligence Unit website.

To assess countries' trade barriers, we make use of data on two types of bilateral trade costs: transport costs and tariffs. We measure country pairs' bilateral transport costs using the U.S. import data, which record both the customs-insurance-freight (cif) and free-on-board (fob) value for most import flows. Restricting our analysis to the preferred screened sample described above, we define transport costs as  $a_{zt}^c = (cif_{zt}^c - fob_{zt}^c) / fob_{zt}^c$  and we estimate *ad valorem* transport costs per mile across all  $z$  in industry  $s$  in year  $t$  by regressing the relative value spent on customs, insurance and freight on imports from country  $c$  on the distance the exports have travelled,

$$\ln(a_{zt}^c) = \delta_{st} \ln(D^{c,US}) + \beta'_{st} X^{c,US} + \epsilon_{zt}^c, \quad (40)$$

where  $D^{c,US}$  represents the great circle distance in miles between the United States and country  $c$  and  $X^{c,US}$  represents additional controls, including whether country  $c$  shares a common language or border with the United States or was ever a colony of the United States. In the estimations below we set  $a_{st}^{cd}$  equal to  $\exp(\widehat{\delta}_{st} \ln D^{cd} + \widehat{\beta}'_{st} X^{cd})$ .

Tariff information is derived from the Trade Analysis and Information System (TRAINS) Database maintained by the United Nations Conference on Trade and Development (UNCTAD). In principle, these data record countries' most favored nation (MFN) tariffs as well as any preferential (PRF) tariff rates that might be available for a subset of trading partners at the eight-digit Harmonized System level. In practice, product-country coverage in the dataset is very sparse, hampering our ability to control properly for trade policy in equation (37).

The construction of the trade cost term  $\tau_{st}^c$  is more challenging because it requires values for the unobservable consumption price indexes  $G_s^c$  (equation 27). As indicated in

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<sup>29</sup>Unfortunately, country pairs' reported trade flows with each other are often mutually inconsistent. Since our principal interest is the accuracy of countries' overall net trade with the world, we favor this approach, which maximizes reporting consistency within countries, to the one taken by Feenstra et al. (1997, 2000), which generally relies on reporting countries' import statistics to estimate bilateral trade flows. Further details of our data refinement procedures are described in the separate Data Appendix.

equation (23), up to a factor of proportionality (captured by the constant in the regression), the term  $\sum_z n_z^{c''} (\tilde{p}_z^{c''})^{1-\sigma_s}$  is the share of country  $c''$  in world production of sector  $s$  in a world equilibrium with no trade costs. We approximate this share by the share of country  $c''$  in “world” exports of that sector, i.e., the total exports of all countries in the preferred estimation sample. While this approximation is imperfect, the theoretical and observed shares should both largely be driven by country size. As a result, this approximate measure should capture a substantial fraction of the relevant variation in the unobserved shares. The consumption indexes  $G_s^c$  also require an estimate of the elasticity of substitution  $\sigma_s$ . We compute  $\tau_{st}^c$  using  $\sigma_s = 6$  and note that alternative values of  $\sigma_s$  ranging from 3 to 10 have almost no impact on our results. We obtain the main input for the construction of the trade cost term, the bilateral trade costs  $\tau_s^{cc'}$ , by adding the measures of bilateral transport costs and tariffs explained above.

Finally, to compute countries’ real exchange rates, we use the real effective exchange rate series reported by the Economist Intelligence Unit (EIU) on their website. Though the EIU dataset is reasonably complete, we fill in any holes in it by using data from the World Bank and the International Monetary Fund.

### 5.2.2. Second-Stage Results

Table 4 reports second-stage estimates of  $\gamma_s$  and  $b_s$  from the estimation of equation (37) by OLS and two-stage least squares (2SLS).<sup>30</sup> Robust standard errors adjusted for clustering at the country level are reported below each coefficient. As indicated in the table, the OLS estimate of  $\gamma_s$  has the expected negative sign but is statistically insignificant. The 2SLS estimate, on the other hand, is both negative and statistically significant as well as substantially lower than the OLS estimate, -0.325 versus -0.029. The coefficient on the trade cost term is negative and insignificant for OLS. For 2SLS, it has the expected positive sign but remains statistically insignificant, an outcome that may reflect the difficulties associated with accurate measurement of trade costs. The final row of the table reports an F-statistic for the first stage of 2SLS of 37.8.<sup>31</sup>

Log Quality Index intercepts and slopes, normalized by annual means as in equation (34), are displayed in Figure 3 along with their 95 percent confidence bands.<sup>32</sup> Estimated

<sup>30</sup>Given our rejection of a unit root using the test developed by Levin et al. (2002), we perform the estimation in levels rather than in differences. The test is performed on the dependent variable, each of the regressors, and the residual allowing alternatively for a constant and for both a constant and a time trend. The null hypothesis that there is a unit root is rejected at the 1% significance level in all cases.

<sup>31</sup>By comparison, implementation of the second stage using first-stage estimates based on the “V”, “sink” and HK indexes discussed above yields first-stage F-statistics of 31.9, 34.8, and 3.9, respectively.

<sup>32</sup>Standard errors are computed using the delta method. Intercept and slope coefficients and standard

intercepts are equivalent to countries' relative log export quality in 1989. As indicated in the figure, China's quality in 1989 is two-thirds ( $e^{0.418}$ ) that of the mean country in that year, while Germany's is more than twice as high ( $e^{0.723}$ ). Estimated slopes report how much relative export quality rises or falls *vis a vis* the mean country each year. The countries with the highest and lowest slopes are Ireland and Hong Kong, at 0.076 and -0.037, respectively, though the former has a relatively large confidence interval.

Figure 3 sorts countries according to their intercepts, from low to high. Though these intercepts vary widely, they tend to be high for developed economies like Switzerland and Sweden and low for developing countries like Malaysia and the Philippines. Quality slopes also vary substantially across countries but appear to be inversely related to intercepts. Key outliers to this pattern include Hungary, Singapore and Ireland, though the estimated slopes of the latter two countries are estimated relatively imprecisely.

Normalized Quality Indexes across the sample period are displayed along with 95 percent confidence bands for a set of nine countries in Figure 4. The relative tightness of the confidence bands indicates substantial and statistically significant variation in product quality across both countries and years. China's export quality relative to the mean country in the sample, for example, is substantially and significantly below that of Germany, Japan and even Singapore in all years and is essentially flat over time. Quality indexes for Hungary and Singapore, on the other hand, rise significantly over time, though the trend for Singapore is relatively imprecise.

To assess the robustness of our results, we performed a number of sensitivity analyses. First, we find that our second-stage regression results are not driven by the inclusion of any particular country, as selectively removing each one country from the sample has little impact on our point estimates. Second, we obtain similar results using either more or less stringent secondary screens, though standard errors are generally larger in the latter case. Finally, our second-stage results are very similar when using either the "V" or "sink" estimators discussed above to derive first-stage Impure Price Indexes. Though we do not report all of these results to conserve space, they are available from the authors upon request.

## 6. Evolution of Overall Manufacturing Quality Over Time

This section examines the evolution of countries' relative manufacturing quality over time. We show that several developing countries, notably the Philippines and Malaysia, errors are reported for each country in the separate Data Appendix.

exhibit large increases in relative quality over the sample period, and that changes in countries' raw relative export prices can be a poor approximation of changes in their relative product quality. We also demonstrate that the "quality" gap between initially high- and low-income countries narrows more substantially over time than their "income" gap. Though these findings are primarily descriptive, they nevertheless highlight the range of issues that our estimates might help address.

Table 5 displays countries' overall manufacturing quality rankings at four-year intervals from 1989 to 2003. Countries are sorted according to their ranking in 1989, and the final column of the table reports the change in ranking over the sample period. Countries whose rank changes by more than ten places between 1989 and 2003 are highlighted.

As indicated in the table, quality rankings are generally more stable for countries that begin the sample period with relatively high normalized Quality Indexes. Among the upper tercile of countries (from Switzerland to Norway), the average change in rank is -1.7. In this group, Japan and Spain exhibit the largest declines, from 10 to 17 and from 14 to 21, respectively, while Ireland experiences the largest gain, from 7 to 1.

More substantial re-rankings are observed among countries that begin the sample period with low quality. Across countries in the bottom tercile (Turkey to Chile), for example, the average change in rank is 2.6. The largest increases in this cohort are exhibited by Malaysia and the Philippines, which jump 27 and 18 places, respectively, between 1989 and 2003. The sharpest declines of this group are registered by Colombia, China and India at -11, -9 and -9, respectively. Across all countries, Singapore registers the second highest jump in ranking, from 23 in 1989 to 2 in 2003.

Changes in quality rankings inferred from our method can be quite different from changes in countries relative export prices, indicating that naively equating price and quality can yield misleading results. Figure 5 compares the change in countries' normalized log Quality Indexes versus their change in normalized log Impure Price Indexes between 1989 and 2003. Though these two changes are positively correlated (0.32), quality and prices move in opposite directions for one third of the sample. For Malaysia, Singapore, Thailand and Indonesia, quality rises while raw export prices fall, while the opposite is observed for Argentina, Greece, Israel, Mexico, Poland, Portugal, Spain and Switzerland. These divergences between quality and impure prices are due to variation in countries' global net trade balances. As illustrated in Figure 6, for example, Malaysia's rising trade balance combined with its relatively flat impure prices results in rising estimated quality.

The attainment of higher quality levels in manufacturing is often thought to foster export development, which in turn induces economic growth. Seminal research by Gross-

man and Helpman (1991) and Aghion and Howitt (1992), for example, highlights a link between countries' ability to "climb the quality ladder" and their economic development. In these models, productivity gains allow countries to produce higher quality goods and thereby achieve higher income. Based on similar reasoning, governments in several countries have implemented export development programs that help firms improve the quality of their products. Despite its importance, there has been relatively little empirical investigation into the link between product quality and economic development, mostly due to a lack of information about product quality.

Here, we find a strong positive relationship between countries' manufacturing quality and income, with the cross-country correlation of normalized log manufacturing quality and normalized log income per capita ranging from a high of 0.88 in 1990 to a low of 0.66 in 2003.<sup>33</sup> The correlation between *growth* in manufacturing quality and *growth* in per capita income between 1989 and 2003, however, is substantially smaller at 0.30. This relationship is displayed in Figure 7. Prominent outliers to the positive correlation include the Philippines and China. The Philippines saw its quality rise relative to the mean country while its income fell. The opposite is true for China, whose quality fell 11.1 percent relative to the average country even as its per capita income grew 130.2 percent.

Finally, we find that trends in quality and income per capita growth are quite different for countries that begin the sample period with either high or low levels of income per capita. Dividing the countries in our sample into two groups according to whether their 1989 income per capita is above or below the median in that year, we find that the 1989 gap in mean quality levels between these two groups is initially lower than the mean income gap, at 0.83 log points versus 2.32 log points, respectively. Over time, as reported in Figure 8, we find substantially stronger narrowing of mean quality levels versus mean income levels. While the quality gap narrows by nearly 40 percent, to 0.51 log points in 2003, the income per capita gap falls only slightly, to 2.25 log points.

## 7. Estimating Export Quality Within Manufacturing

As noted previously, our method for identifying product quality contains an aggregation trade-off. While product quality is more likely to be constant across more finely disaggregated products, data on countries' global net trade and trade barriers is scarcer

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<sup>33</sup>Even though Hummels and Klenow (2005) cannot obtain quality levels for each country, they are able to compute estimates of the cross-sectional elasticity of quality with respect to income using assumptions about the elasticity of substitution and the number of varieties countries produce. Depending on these assumptions, their estimate of the cross-sectional quality-income elasticity in 1995 ranges from 0.09 to 0.23.

and potentially more susceptible to measurement error. In this section we compute Quality Indexes across the four one-digit SITC industries that constitute manufacturing. We also estimate quality across the two two-digit SITC sectors that represent apparel and textiles, respectively, in order to explore the potential influence of countries' use of intermediate inputs outside of the sectors at which quality is being estimated.

### 7.1. *One-digit SITC Sectors*

Estimation of export quality within manufacturing relies upon the same strategies and datasets described above. To conserve space, we omit a discussion of screening, but note that primary and secondary screens exert similar influence. The number of countries that can be included in the analysis varies by industry because all countries do not participate equally in each industry. Of the 43 countries used for aggregate manufacturing, we are left with 27 countries in Chemicals, 41 countries in Manufactured Materials, 37 countries in Machinery and 41 countries in Miscellaneous Manufacturing.

Table 6 reports estimation results as well as details of the first-stage estimation sample by industry and year. Across industries, the data are thicker in terms of product-country-pair observations and median products in common for Manufactured Materials and Miscellaneous Manufactures than for Machinery and Chemicals. Goodness of fit in terms of the share of estimates falling within bounds is highest in Machinery and lowest in Manufactured Materials and Miscellaneous Manufactures, and this ordering generally remains consistent with the ordering of their Paasche-Laspeyres intervals from high to low.

Figure 9 reports estimates of countries' normalized first-stage Impure Price Indexes by manufacturing industry for 2003 versus 1989. As indicated in the figure, prices are most tightly distributed in Chemicals (except for outlier Ireland) and are most dispersed in Miscellaneous Manufactures. We find that countries' Impure Price Indexes are positively correlated across industries. In 2003, this correlation is highest for Manufactured Materials versus Miscellaneous Manufactures (0.82) and lowest for Chemicals versus Machinery (0.54).

Table 7 reports the second-stage OLS (top panel) and 2SLS (bottom panel) estimates of  $\gamma_s$  and  $b_s$  by industry. The 2SLS estimates of  $\gamma_s$  have the expected negative sign in all four industries, but the estimate for Chemicals is statistically insignificant. As with the results for aggregate manufacturing above, the 2SLS coefficients for the trade cost term have the expected positive sign but are statistically insignificant. The strength of the real exchange rate as an instrument for the Impure Price Index varies across industries.

The F-statistic for the first stage of the 2SLS regression is high for both Machinery and Miscellaneous Manufactures, low for Manufactured Materials, but especially low (0.01) for Chemicals. A potential explanation for this result is that Chemical products are less horizontally differentiated than products in Machinery or Miscellaneous Manufactures. If that were the case, export prices might not be responsive to movements in countries' real exchange rate and instead respond mostly to movements in world prices. This hypothesis receives some support from the relatively low price dispersion exhibited in the Chemical Impure Price Indexes (Figure 9).

Normalized log Quality Index intercepts and slopes along with their standard errors are displayed along with their 95 percent confidence bands in Figure 10.<sup>34</sup> Outside of Chemicals, most countries' normalized slopes and intercepts are significantly different from zero (i.e., the mean country). As with aggregate manufacturing, the ordering of quality intercepts generally accords with expectations: developed economies have the highest intercepts in Machinery, for example, while Italy's intercepts are among the highest in Manufactured Materials and Miscellaneous Manufactures, which include Textiles (SITC 65) and Apparel (SITC 84), respectively. China's quality ranking, on the other hand, is relatively low in both Manufactured Materials and Machinery and relatively high in Miscellaneous Manufacturing. Given the weak results for the Chemicals sector, we exclude it from further analysis.

Disaggregated quality estimates reveal substantial variation in export quality across industries within countries. While quality intercepts for Manufactured Materials and Miscellaneous Manufactures are positively correlated, the correlation of intercepts for Machinery and Miscellaneous Manufactures is approximately zero. Hong Kong and Taiwan, for example, have relatively high intercepts for Miscellaneous Manufactures but are in the middle of the pack for Machinery. Quality Index slopes display similar variation: across countries the non-Chemical industry slopes have the same sign in only 17 of the 43 countries in the sample. These differences are highlighted in Figure 11 for the subset of nine countries examined above. For Singapore, quality increases relatively strongly compared to the mean country in all three sectors. For Malaysia, quality increases strongly in Machinery, much more moderately in Manufactured Materials, and declines in Miscellaneous Manufactures.

Quality convergence within manufacturing also varies across industries. Figure 12 reports the evolution of mean quality for countries with initially high and low per capita

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<sup>34</sup>As noted above, standard errors are computed using the delta method. Intercept and slope coefficients and standard errors are reported for each country and industry in the separate Data Appendix.

income. As indicated in the figure, there is a substantial narrowing of quality in Machinery, weaker convergence in Miscellaneous Manufactures and an unchanging quality gap in Manufactured Materials.

### 7.2. *Two-digit SITC Sectors: Apparel and Textiles*

As noted in the introduction, our method for estimating product quality involves an aggregation trade-off. For broad SITC categories such as all manufacturing, the assumption of constant quality across all products in the category is strong, but data on countries' global net trade is more readily available and more likely to be reliable. Another potential advantage of using broader SITC sectors is their ability to dampen the impact of countries' use of imported intermediate inputs. Use of such inputs is an issue when they straddle the sectors at which quality is being estimated. Countries with a strong comparative advantage in one sector, for example, might be large net exporters of that sector but large net importers of a second sector which is an input to the first. All else equal, this situation may lead quality in the first and second sectors to be over- and underestimated, respectively. In principle this problem can be solved by using input-output tables to map imported intermediates to final goods. Unfortunately, such tables are unavailable at a useful level of aggregation for a wide sample of countries. Another potential solution would be to use value added trade data, but these data, too, are generally unavailable.

In this section we investigate the potential impact of input-output linkages across sectors by examining the relationship between textiles (SITC 65) and apparel (SITC 84). We also explore a potential solution to the intermediate-inputs problem that involves combining linked industries into a more aggregate sector. First, we estimate quality for textiles and apparel separately. We then estimate quality for the hybrid "Apparel & Textiles", which includes products from both industries. We find that even though cross-country correlations of apparel and Apparel & Textile quality are high in every year, estimated quality for apparel declines substantially for some countries when textile deficits are taken into account.

Across the sample period, an average trade surplus in apparel coincides with an average trade deficit in textiles in eight developing countries: Colombia, Hong Kong, Israel, Mexico, Malaysia, the Philippines, Portugal and South Africa. The contrast is most stark for Hong Kong, whose average net trade in SITC 65 and SITC 84 over the sample period is -1.0 and 7.6 percent of GDP, respectively.<sup>35</sup> The opposite pattern, that is, an average

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<sup>35</sup>Interestingly, Hong Kong's net trade in these two sectors move in opposite directions over time, with the result that its declining trade surplus in Apparel & Textiles is less extreme than its declining trade

deficit in apparel and an average surplus in textiles, is exhibited by several developed economies: Austria, Belgium, France, Germany, Japan and Switzerland. The remaining countries either have a surplus in both sectors (e.g., Italy, Pakistan and Romania), or a deficit in both sectors (e.g., Spain, Ireland and Sweden). The largest re-ranking of apparel versus Apparel & Textile qualities occurs for Pakistan, which has a strong surplus in both apparel and textiles.

Table 8 reports 2SLS estimates of  $\gamma_s$  and  $b_s$  for SITC 65, SITC 84 and the hybrid Apparel & Textiles. The estimates of  $\gamma_s$  for all three remain negative and significant, while the estimates for the trade cost term have the indicated signs but remain insignificant. Though the Quality Index intercepts and slopes for the hybrid sector versus those for SITC 84 differ markedly for some countries, their correlations across all countries are quite high: 0.90 for the intercepts and 0.91 for the slopes. Estimated Quality Indexes are also highly correlated across years, ranging from 0.90 in 1989 to 0.74 in 2003.

These findings suggest that even though controlling for intermediate inputs does not appear to change the overall pattern of results across countries, it can have a substantial effect on individual countries' estimated quality. As a result, it would be prudent to include as much information about input-output linkages as possible when estimating quality at low levels of aggregation. Over time, collection and dissemination of more detailed data on countries' international trade and use of inputs should make this task easier.

## 8. Conclusion

This paper attempts to fill an important gap in the international trade and development literature by outlining a method for identifying how countries' product quality evolves over time. First, we show how an important but unobserved Impure Price Index comparing countries' export prices is bounded by their observable Paasche and Laspeyres indexes. Then, we develop a method for decomposing an estimate of this Impure Price Index into Quality and Pure Price Indexes. This method makes use of information on consumers' valuation of countries' products contained in their net trade with the world and allows for both vertical and horizontal product differentiation. In contrast to a vast literature that associates cross-country variation in export unit-values with variation in product quality – implicitly assuming away cross-country variation in quality-adjusted prices – our method allows for price variation induced by factors other than quality, e.g.,

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surplus in apparel. Between 1989 and 2003, Hong Kong's apparel surplus declines from 13.4 to 5.3 percent of GDP, while its Apparel & Textiles surplus declines from 11.2 to 5.6 percent of GDP.

comparative advantage or currency misalignment.

Implementation of our method reveals trends in product quality not evident in export prices alone. Indeed, for several countries, export prices and quality evolve quite differently. Our estimation also highlights the importance of accounting for intermediate trade in estimating countries' export quality. Further theoretical and empirical efforts on this front will be quite useful.

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## A Theory Appendix

### A1. Proof of Proposition 1

We have already shown that  $\ln H_s^{cd} \leq \ln P_s^{cd} + \ln \phi_s^c$ . To demonstrate that  $\ln H_s^{cd} \leq \ln P_s^{cd}$ , we need to show that  $\ln \phi_s^c \leq 0$ .

Using the fact that  $n_z^c = \frac{\bar{n}_z^c}{\bar{n}_s^o} (\tilde{n}_z^{c,cd} + \hat{n}_z^{cd})$  and noting that the set  $I_s^{cd}$  results from subtracting  $U_s^{cd}$  from  $I_s$ , we can write:

$$\sum_{z \in I_s^{cd}} n_z^c \Delta \hat{p}_z^{cd} = \frac{\bar{n}_s^c}{\bar{n}_s^o} \left[ \sum_{z \in I_s^{cd}} \tilde{n}_z^{c,cd} \Delta \hat{p}_z^{cd} + \sum_{z \in I_s} \hat{n}_z^{cd} \Delta \hat{p}_z^{cd} - \sum_{z \in U_s^{cd}} \hat{n}_z^{cd} \Delta \hat{p}_z^{cd} \right]. \quad (41)$$

Using the definition of sample covariance provided in section 3, for two variables  $x_z$  and  $y_z$  it is true that

$$\sum_{z \in A} x_z y_z = Z_A \text{COV}_A(x_z, y_z) + \frac{1}{Z_A} \sum_{z \in A} x_z \sum_{z \in A} y_z. \quad (42)$$

Then, the first term on the right hand side of (41) can be expressed as

$$\begin{aligned} \sum_{z \in I_s^{cd}} \tilde{n}_z^{c,cd} \Delta \hat{p}_z^{cd} &= Z_s^{cd} \text{COV}_{I_s^{cd}}(\tilde{n}_z^{c,cd}, \Delta \hat{p}_z^{cd}) + \frac{1}{Z_s^{cd}} \sum_{z \in I_s^{cd}} \tilde{n}_z^{c,cd} \sum_{z \in I_s^{cd}} \Delta \hat{p}_z^{cd} \\ &= Z_s^{cd} \text{COV}_{I_s^{cd}}(\tilde{n}_z^{c,cd}, \Delta \hat{p}_z^{cd}) - \frac{1}{Z_s^{cd}} \sum_{z \in U_s^{cd}} \tilde{n}_z^{c,cd} \sum_{z \in I_s^{cd}} \Delta \hat{p}_z^{cd} \end{aligned} \quad (43)$$

where the second equality uses the property  $\sum_{z \in I_s} \tilde{n}_z^{c,cd} = \sum_{z \in I_s^{cd}} \tilde{n}_z^{c,cd} + \sum_{z \in U_s^{cd}} \tilde{n}_z^{c,cd} = 0$ . In turn, since  $\hat{n}_z^{cd} = \tilde{n}_z^{cd} + \bar{n}_z$ , the second term on the right hand side of (41) can be written as

$$\begin{aligned} \sum_{z \in I_s} \hat{n}_z^{cd} \Delta \hat{p}_z^{cd} &= \sum_{z \in I_s} \tilde{n}_z^{cd} \Delta \hat{p}_z^{cd} + \sum_{z \in I_s} \bar{n}_z \Delta \hat{p}_z^{cd} \\ &= Z_s \text{COV}_{I_s}(\tilde{n}_z^{cd}, \Delta \hat{p}_z^{cd}) + \sum_{z \in I_s} \bar{n}_z \Delta \hat{p}_z^{cd} = 0 \end{aligned} \quad (44)$$

where the first equality in the second line uses (42) and  $\sum_{z \in I_s} \tilde{n}_z^{cd} = 0$ , and the second equality uses Assumption 4 and the definition of  $\tilde{P}_s^{cd}$  in (7), which implies that  $\sum_{z \in I_s} \bar{n}_z \Delta \hat{p}_z^{cd} = 0$ .

Combining the results in (43) and (44), we can rewrite equation (41) as

$$\begin{aligned} \sum_{z \in I_s^{cd}} n_z^c \Delta \tilde{p}_z^{cd} &= \frac{\bar{n}_s^c}{\bar{n}_s^o} \left[ Z_s^{cd} \text{cov}_{I_s^{cd}}(\tilde{n}_z^{c,cd}, \Delta \tilde{p}_z^{cd}) - \sum_{z \in U_s^{cd}} \tilde{n}_z^{c,cd} \frac{\sum_{z \in I_s^{cd}} \Delta \tilde{p}_z^{cd}}{Z_s^{cd}} - \sum_{z \in U_s^{cd}} \hat{n}_z^{cd} \Delta \tilde{p}_z^{cd} \right] \\ &\geq -\frac{\bar{n}_s^c}{\bar{n}_s^o} \left[ \sum_{z \in U_s^{cd}} \left( \tilde{n}_z^{c,cd} \frac{\sum_{z \in I_s^{cd}} \Delta \tilde{p}_z^{cd}}{Z_s^{cd}} + \hat{n}_z^{cd} \Delta \tilde{p}_z^{cd} \right) \right] = 0, \end{aligned}$$

where the inequality uses Assumption 3 and the last equality uses Assumption 5.

Finally, decomposing  $\Delta \tilde{p}_z^{cd}$  according to its definition in (21), we obtain

$$\ln \phi_s^c \equiv \ln \left( \frac{\sum_{z \in I_s^{cd}} n_z^c \left( \frac{\tilde{p}_z^c}{P_s^c} \right)^{1-\sigma_s}}{\sum_{z \in I_s^{cd}} n_z^c \left( \frac{\tilde{p}_z^d}{P_s^d} \right)^{1-\sigma_s}} \right)^{\frac{1}{1-\sigma_s}} \leq 0. \quad (45)$$

Substituting this result into  $\ln H_s^{cd} \leq \ln P_s^{cd} + \ln \phi_s^c$  in equation (15), we obtain

$$\ln H_s^{cd} \leq \ln P_s^{cd}. \quad (46)$$

An analogous proof shows that  $\ln P_s^{cd} \leq \ln L_s^{cd}$ . Hence, the Paasche and Laspeyres indexes bound the Impure Price Index,

$$\ln H_s^{cd} \leq \ln P_s^{cd} \leq \ln L_s^{cd}.$$

## A2. Proof of Proposition 2

Solving for  $n_z^c$  in equation (4) and substituting the result into (25), we can rewrite the right-hand side of that equation as

$$-1 + \frac{\bar{n}_s^c}{\bar{n}_s^o} \frac{1}{Y^c} \left( \sum_{c'} Y^{c'} \left( \frac{\tau_s^{cc'}}{G_s^{c'}} \right)^{1-\sigma_s} \right) \left( \sum_z \bar{n}_z (\tilde{p}_z^c)^{1-\sigma_s} + \left( \tilde{P}_s^c \right)^{1-\sigma_s} \sum_z \tilde{n}_z^c \left( \frac{\tilde{p}_z^c}{\tilde{P}_s^c} \right)^{1-\sigma_s} \right).$$

Using the definition of  $\tilde{P}_s^c$  in equation (7) and the fact that, since  $\sum_z \tilde{n}_z^c = 0$ ,  $\sum_z \tilde{n}_z^c \left( \frac{\tilde{p}_z^c}{\tilde{P}_s^c} \right)^{1-\sigma_s} = Z_s \text{cov} \left[ \tilde{n}_z^c, \left( \frac{\tilde{p}_z^c}{\tilde{P}_s^c} \right)^{1-\sigma_s} \right]$ , the above expression can be further rewritten as

$$-1 + \frac{\bar{n}_s^c}{\bar{n}_s^o} \frac{1}{Y^c} \left( \tilde{P}_s^c \right)^{1-\sigma_s} \left( \sum_{c'} Y^{c'} \left( \frac{\tau_s^{cc'}}{G_s^{c'}} \right)^{1-\sigma_s} \right) \left( 1 + Z_s \text{cov} \left[ \tilde{n}_z^c, \left( \frac{\tilde{p}_z^c}{\tilde{P}_s^c} \right)^{1-\sigma_s} \right] \right).$$

Using Assumption 6 and equation (26), we can substitute the latter expression for the right-hand side of (25). Rearranging terms and taking natural logarithms, we obtain

$$\ln \left( 1 + \frac{1}{b_s} \frac{T_s^c}{Y^c} \right) = k_s \ln \left[ \left( \tilde{P}_s^{co} \right)^{1-\sigma_s-\eta_s} \left( \sum_{c'} Y^{c'} \left( \frac{\tau_s^{cc'}}{G_s^{c'}} \right)^{1-\sigma_s} \right) [1 + Z_s (V_s + \mu_s^c)] \right] \quad (47)$$

where  $k_s = \ln \left[ \frac{(\tilde{P}_s^o)^{1-\sigma_s}}{Y^o} \right]$ . Using  $\ln(1+x) \simeq x$ , and abstracting from the approximation error, we can express equation (47) as

$$\frac{1}{b_s} \frac{T_s^c}{Y^c} = k_s + (1 - \sigma_s - \eta_s) \ln \tilde{P}_s^{co} + \ln \left( \sum_{c'} Y^{c'} \left( \frac{\tau_s^{cc'}}{G_s^{c'}} \right)^{1-\sigma_s} \right) + Z_s (V_s + \mu_s^c). \quad (48)$$

Finally, we can express this equation as

$$\frac{T_s^c}{Y^c} = \Psi_s + \gamma_s \ln \tilde{P}_s^{co} + b_s \tau_s^c + \iota_s^c$$

where

$$\begin{aligned} \gamma_s &= b_s (1 - \sigma_s - \eta_s) < 0, & \Psi_s &= b_s (k_s + Z_s V_s), \\ \tau_s^c &= \ln \left( \sum_{c'} Y^{c'} \left( \frac{\tau_s^{cc'}}{G_s^{c'}} \right)^{1-\sigma_s} \right), & k_s &= \ln \left[ \frac{(\tilde{P}_s^o)^{1-\sigma_s}}{Y^o} \right], \\ \iota_s^c &= b_s Z_s \mu_s^c \end{aligned}$$

	Percent of Unscreened Sample's Import Value	Explicit Country-Pair Comparisons	Median Common Products	Median Interval Length	Correctly Ordered Bounds
Unscreened Sample	100.0%	861	1213	4.46	99.9%
Primary Screened Sample	99.8%	861	347	2.51	99.8%
Preferred Secondary Screened Sample	88.8%	829	228	0.74	99.4%

Notes: Table displays several attributes of the estimation sample for all manufacturing in 2003 according to three methods of screening the raw data. All samples contain the same set of 43 countries. Import value for each sample is expressed as a percentage of the unscreened sample. Explicit country-pair comparisons is the number of country pairs that appear in the sample; the maximum is 903 (i.e.,  $43 \times 42 / 2$ ). Median common products is the median number of export products exported in common to the United States across country pairs appearing in the sample. Median interval length is the median log difference between Paasche and Laspeyres bounds. Correctly ordered bounds is the percent of bounds in the sample for which the Paasche index is less than the Laspeyres index.

Table 1: Sample Attributes for All Manufacturing in 2003, by Screening

Attributes of Estimation Sample						Attributes of First-Stage Estimation		
Countries	Country-Pair		Product-Country-Pair Observations	Median Common Products Across Country Pairs	Median Log Paasche- Laspeyres Interval	Objective Function	Maximization Standard Error	First-Stage Estimates Within Bounds
	Observations	Country Pairs						
1989	43	811	133	208,108	0.74	-349	0.16	90.4%
1990	43	829	143	223,564	0.68	-334	0.14	90.8%
1991	43	814	144	219,596	0.69	-332	0.15	91.5%
1992	43	814	146	224,875	0.73	-322	0.14	91.2%
1993	43	823	156	239,190	0.74	-319	0.16	90.6%
1994	43	846	171	263,986	0.73	-320	0.16	91.8%
1995	43	858	185	292,615	0.76	-272	0.14	94.2%
1996	43	862	190	308,028	0.72	-251	0.13	93.5%
1997	43	866	206	328,629	0.69	-310	0.15	93.3%
1998	43	869	221	342,476	0.73	-291	0.15	93.5%
1999	43	873	226	350,882	0.76	-245	0.14	93.7%
2000	43	877	249	374,151	0.72	-300	0.16	93.0%
2001	43	875	238	358,160	0.78	-222	0.14	94.1%
2002	43	831	234	341,940	0.74	-239	0.15	94.5%
2003	43	829	228	330,968	0.74	-271	0.16	93.8%

Notes: First panel displays characteristics of the preferred first-stage estimation sample, by year. Second panel displays attributes of the first-stage estimation.

Table 2: Sample and First-Stage Estimation Attributes, All Manufacturing

	First-Stage Estimator			
	Preferred	V	Sink	HK
1989	90.4	88.2	86.1	46.4
1990	90.8	89.3	85.3	37.0
1991	91.5	90.0	84.8	36.7
1992	91.2	90.7	85.9	37.5
1993	90.6	89.4	86.9	37.6
1994	91.8	90.5	89.0	40.2
1995	94.2	91.7	88.5	45.4
1996	93.5	93.1	90.2	46.8
1997	93.3	93.5	87.8	46.8
1998	93.5	94.2	88.1	43.8
1999	93.7	94.0	87.9	44.7
2000	93.0	92.2	88.3	47.0
2001	94.1	94.2	89.9	44.0
2002	94.5	93.4	86.8	41.7
2003	93.8	92.7	90.5	45.3
Mean	92.7	91.8	87.7	42.7

Notes: Table compares the share of first-stage estimates lying between country-pairs' Paasche and Laspeyres bounds, by year.

Table 3: Goodness of Fit Across Alternative First-Stage Estimators, By Year

	OLS	2SLS
Impure Price Index	-0.029	-0.325 ***
	0.020	0.094
Trade Costs	-0.016	0.022
	0.062	0.080
Observations	640	640
R <sup>2</sup>	0.93	0.87
First-Stage F Statistic	.	37.8

Notes: Table displays OLS and 2SLS regression results for estimation of equation (37) on the preferred sample (see text). Coefficients for country fixed effects and time trends are omitted. Heteroskedasticity-robust standard errors adjusted for clustering at the country level are reported below each coefficient. The instrument for the Impure Price Index in the 2SLS results is the real exchange rate. \*\*\* denotes statistical significance at the 1 percent level.

Table 4: Second-Stage Regression Results for All Manufacturing

Country	1989	1993	1998	2003	Change
Switzerland (CHE)	1	2	2	3	-2
Sweden (SWE)	2	3	3	5	-3
Germany (DEU)	3	4	6	8	-5
France (FRA)	4	6	9	9	-5
Italy (ITA)	5	7	8	7	-2
Finland (FIN)	6	5	4	4	2
Ireland (IRL)	7	1	1	1	6
Belgium (BEL)	8	8	5	6	2
Denmark (DNK)	9	9	10	11	-2
Japan (JPN)	10	10	12	17	-7
United Kingdom (GBR)	11	11	13	16	-5
Austria (AUT)	12	12	11	15	-3
Netherlands (NLD)	13	13	15	12	1
Spain (ESP)	14	17	19	21	-7
Norway (NOR)	15	14	14	10	5
Australia (AUS)	16	18	22	27	-11
Israel (ISR)	17	16	16	19	-2
Hong Kong (HKG)	18	22	26	33	-15
Canada (CAN)	19	21	23	25	-6
Taiwan (TWN)	20	20	21	24	-4
New Zealand (NZL)	21	24	27	32	-11
Korea, Republic of (KOR)	22	19	17	18	4
Singapore (SGP)	23	15	7	2	21
Portugal (PRT)	24	23	20	20	4
Argentina (ARG)	25	26	25	23	2
Brazil (BRA)	26	27	28	30	-4
Hungary (HUN)	27	25	18	13	14
Mexico (MEX)	28	28	33	34	-6
Turkey (TUR)	29	29	32	31	-2
Colombia (COL)	30	31	36	41	-11
India (IND)	31	33	37	40	-9
South Africa (ZAF)	32	30	31	29	3
China (CHN)	33	36	40	42	-9
Greece (GRC)	34	35	35	37	-3
Poland (POL)	35	37	38	39	-4
Indonesia (IDN)	36	34	30	26	10
Romania (ROM)	37	40	41	38	-1
Morocco (MAR)	38	38	34	28	10
Thailand (THA)	39	41	39	35	4
Philippines (PHL)	40	39	29	22	18
Malaysia (MYS)	41	32	24	14	27
Pakistan (PAK)	42	43	43	43	-1
Chile (CHL)	43	42	42	36	7

Notes: Table records countries' quality ranking according to their normalized quality indexes in each year. Countries are sorted according to their 1989 ranking. Final column reports change between 1989 and 2003. Countries whose rank changes by more than ten places are highlighted.

Table 5: Normalized Quality Rankings

	First-Stage Estimates				Product-Country Pairs	Country Pair Obs	Median Common Products Across Country Pairs	First-Stage Estimates				Product-Country Pairs	Country Pair Obs	Median Common Products Across Country Pairs
	Within Bounds	Objective Function	Maximization Standard Error	Median Log Paasche-Laspeyres Interval				Within Bounds	Objective Function	Maximization Standard Error	Median Log Paasche-Laspeyres Interval			
Chemicals (SITC 5)							Manufactured Materials (SITC 6)							
1989	94.7%	-44	0.08	0.64	16,042	176	56	90.3%	-189	0.13	0.58	59,398	499	66
1990	93.4%	-40	0.08	0.69	18,085	198	59	90.2%	-197	0.12	0.57	61,994	512	68
1991	95.0%	-45	0.11	0.71	17,392	186	63	89.7%	-211	0.13	0.55	59,284	510	66
1992	95.6%	-54	0.11	0.73	20,035	212	62	88.9%	-224	0.15	0.55	61,843	530	67
1993	92.8%	-60	0.12	0.72	21,526	220	65	91.3%	-238	0.15	0.59	67,827	558	72
1994	96.3%	-54	0.11	0.71	23,017	226	66	90.1%	-274	0.16	0.61	76,301	594	76
1995	94.9%	-76	0.13	0.67	25,889	245	67	90.4%	-277	0.17	0.66	84,583	614	82
1996	95.1%	-47	0.07	0.76	27,998	262	70	91.4%	-248	0.16	0.66	88,890	616	84
1997	93.7%	-87	0.12	0.66	30,769	277	72	92.5%	-238	0.15	0.65	94,855	628	88
1998	93.7%	-88	0.13	0.75	32,375	285	71	91.3%	-262	0.16	0.65	101,679	646	91
1999	94.0%	-87	0.14	0.73	32,863	288	72	93.0%	-233	0.14	0.66	104,242	658	94
2000	94.3%	-85	0.12	0.65	34,919	295	74	91.7%	-293	0.18	0.66	111,155	681	96
2001	92.2%	-109	0.15	0.68	34,193	293	74	93.0%	-234	0.15	0.72	103,982	658	94
2002	86.9%	-127	0.21	0.65	33,684	278	73	94.7%	-189	0.12	0.70	101,514	631	95
2003	84.1%	-125	0.21	0.70	33,117	276	75	92.3%	-204	0.15	0.69	93,619	589	94
Machinery (SITC 7)							Miscellaneous Manufactures (SITC 8)							
1989	95.2%	-95	0.14	0.76	43,580	365	77	80.0%	-302	0.12	0.45	76,610	651	71
1990	92.1%	-75	0.11	0.77	44,778	374	78	78.0%	-329	0.13	0.43	86,114	679	79
1991	91.1%	-94	0.12	0.79	46,765	400	77	81.6%	-291	0.13	0.43	82,742	643	76
1992	91.9%	-97	0.12	0.85	44,618	384	76	82.6%	-287	0.13	0.45	85,662	647	79
1993	91.4%	-140	0.17	0.80	47,232	413	75	84.1%	-290	0.14	0.52	89,817	651	79
1994	93.2%	-90	0.10	0.77	54,816	437	84	85.8%	-289	0.14	0.49	96,176	674	83
1995	94.7%	-130	0.13	0.83	66,636	508	90	87.5%	-279	0.13	0.46	101,637	682	91
1996	92.4%	-154	0.13	0.70	74,858	525	96	88.5%	-266	0.14	0.50	102,411	680	88
1997	90.9%	-184	0.15	0.64	80,935	532	100	92.0%	-231	0.11	0.51	107,525	699	94
1998	93.8%	-150	0.12	0.72	82,866	559	97	90.4%	-200	0.11	0.55	111,583	699	98
1999	92.5%	-163	0.14	0.73	85,992	560	100	88.2%	-266	0.14	0.56	113,628	695	102
2000	92.5%	-137	0.12	0.72	93,946	578	107	88.4%	-312	0.16	0.52	120,254	712	110
2001	95.7%	-131	0.13	0.75	91,063	592	102	86.1%	-276	0.14	0.54	113,308	694	107
2002	93.4%	-145	0.15	0.75	86,000	566	98	89.2%	-252	0.15	0.58	106,469	656	102
2003	93.7%	-143	0.13	0.78	84,451	560	98	90.2%	-219	0.12	0.57	105,211	662	102

Notes: Table displays characteristics of the first-stage estimation of Impure Price Indexes by manufacturing industry and year. The number of countries included in the analysis varies by industry: there are 27 in Chemicals, 41 in Manufactured Materials, 37 in Machinery and 41 in Miscellaneous Manufacturing.

Table 6: First-Stage Optimization Results, By Manufacturing Industry

OLS				
	5 - Chemicals	6 - Manuf Mat	7 - Machinery	8 - Misc Manuf
Impure Price Index	0.005 0.005	0.001 0.004	-0.016 0.007	0.001 0.005
Trade Costs	-0.005 0.012	-0.056 0.020	0.016 0.059	-0.001 0.015
Observations	400	608	533	610
R <sup>2</sup>	0.98	0.96	0.92	0.96

2SLS				
	5 - Chemicals	6 - Manuf Mat	7 - Machinery	8 - Misc Manuf
Impure Price Index	-0.100 0.934	-0.202 ** 0.103	-0.110 *** 0.041	-0.054 ** 0.033
Trade Costs	0.109 1.016	0.011 0.041	0.009 0.077	0.016 0.018
Observations	400	608	533	610
R <sup>2</sup>	0.86	0.66	0.86	0.94
First-Stage F Statistic	0.01	4.22	34.3	13.6

Notes: Table reports OLS and 2SLS regression results for equation (37). The instrument for the Impure Price Index is the real exchange rate. Coefficients for country fixed effects and time trends are omitted. Heteroskedasticity-robust standard errors adjusted for clustering at the country level are reported below each coefficient. The instrument for the Impure Price Index in the 2SLS results is the real exchange rate. \*\* and \*\*\* denote statistical significance at the 5 and 1 percent level, respectively.

Table 7: Second-Stage Regression Results, by Manufacturing Industry

	2SLS		
	SITC 65	SITC84	SITC6584
Impure Price Index	-0.023 * 0.015	-0.045 * 0.026	-0.056 * 0.031
Trade Costs	-0.002 0.004	0.002 0.004	-0.003 0.008
Observations	434	528	579
R <sup>2</sup>	0.97	0.93	0.94
First-Stage F Statistic	4.58	11.8	16.0

Notes: Table compares 2SLS regression results for estimation of equation (37) on noted two digit industries and a hybrid industry that combines SITC 65 and SITC 84. The instrument for the Impure Price Index is the real exchange rate. Coefficients for country fixed effects and time trends are omitted. Heteroskedasticity-robust standard errors adjusted for clustering at the country level are reported below each coefficient. \* denotes statistical significance at the 10 percent level.

Table 8: Second-Stage Regression Results for Apparel and Textiles

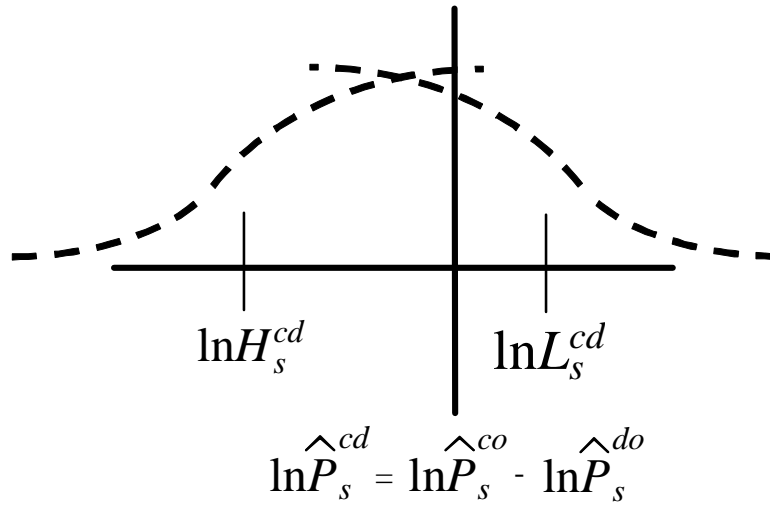


Figure 1: Maximizing the Likelihood that the Observed Paasche and Laspeyres Bounds Contain the Estimated Impure Price Index

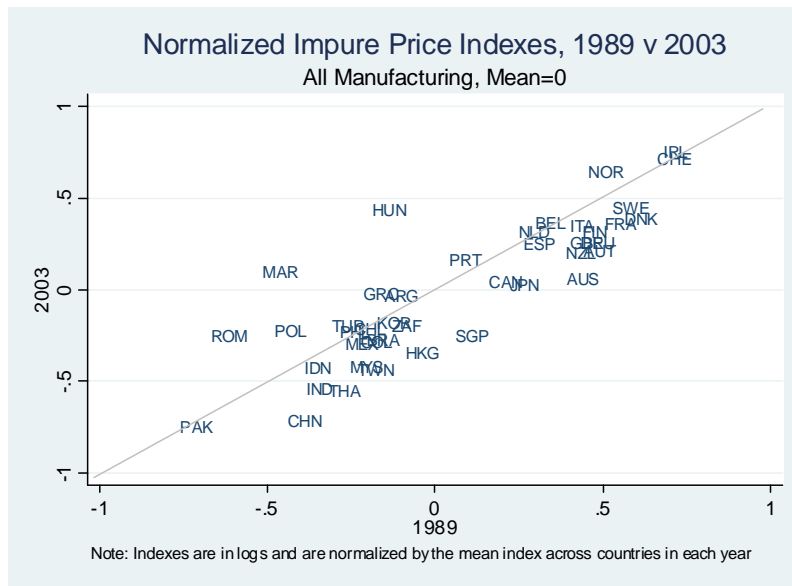


Figure 2: First-Stage Estimated Impure Price Indexes, 2003 versus 1989

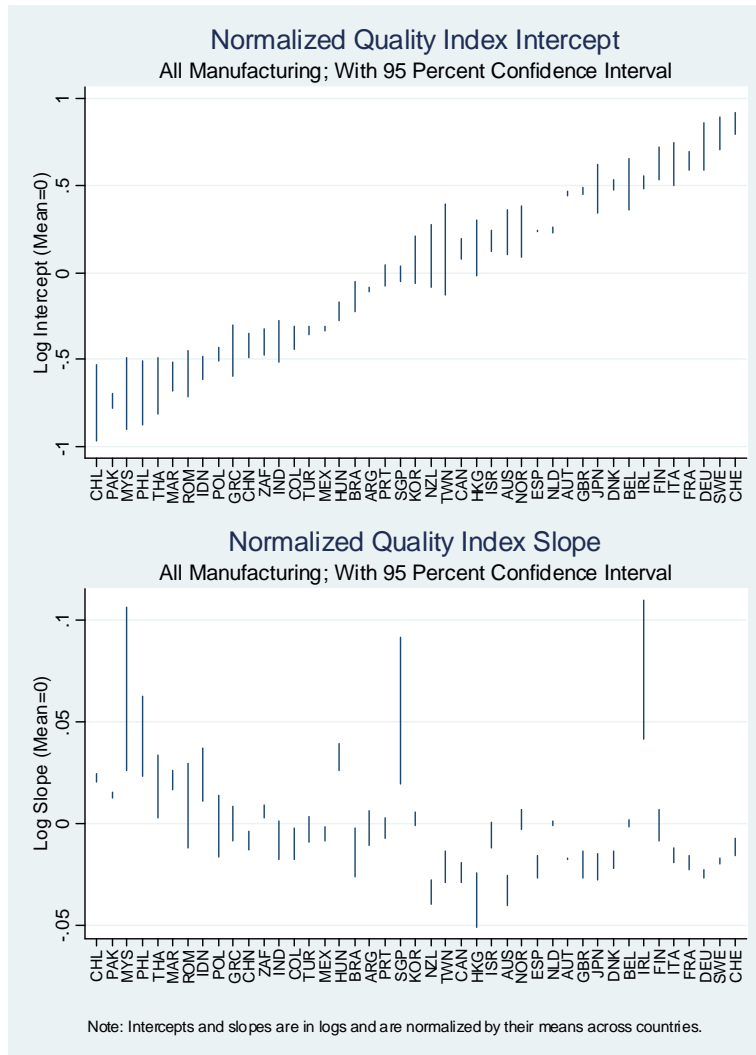


Figure 3: Normalized Log Quality Index Intercept and Slope, by Country

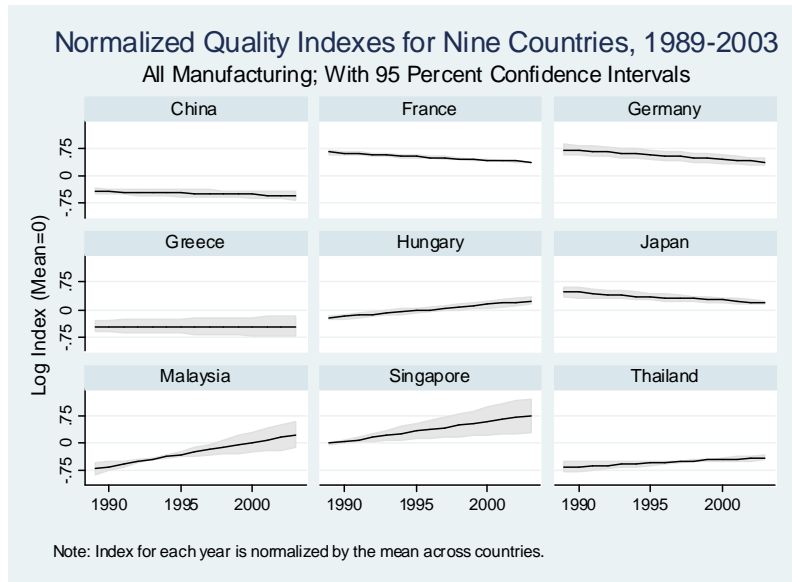


Figure 4: Normalized Log Quality Index for Select Countries, 1989 to 2003

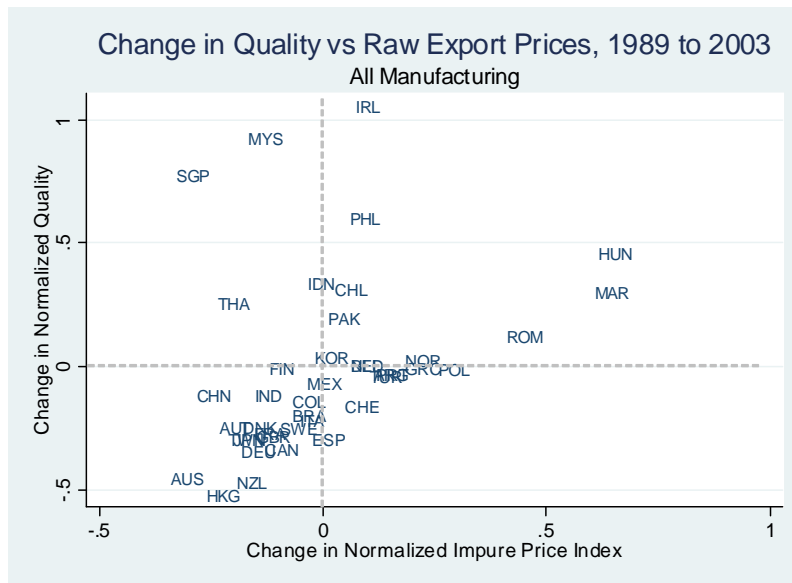


Figure 5: Change in Normalized Quality from 1989 to 2003 vs Change in Normalized Impure Price Indexes 1989 to 2003

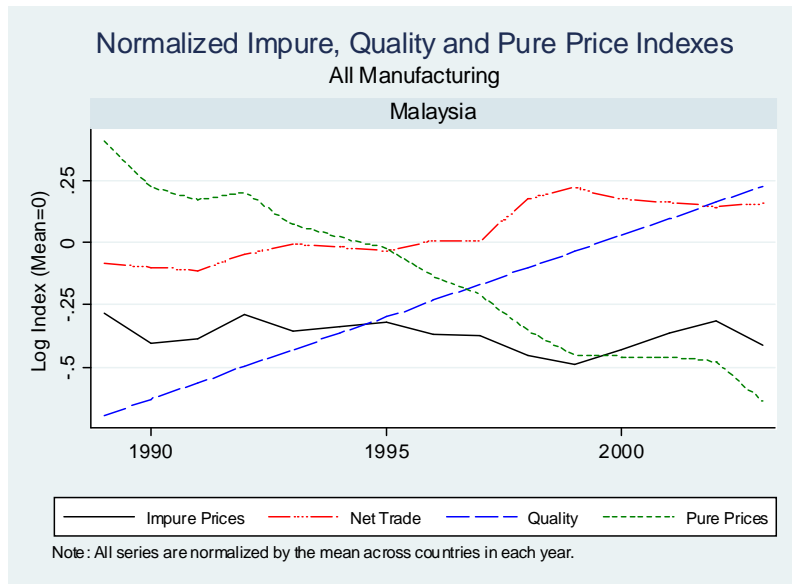


Figure 6: Decomposition of Malaysia’s Impure Price Index into Quality and Pure Price Indexes

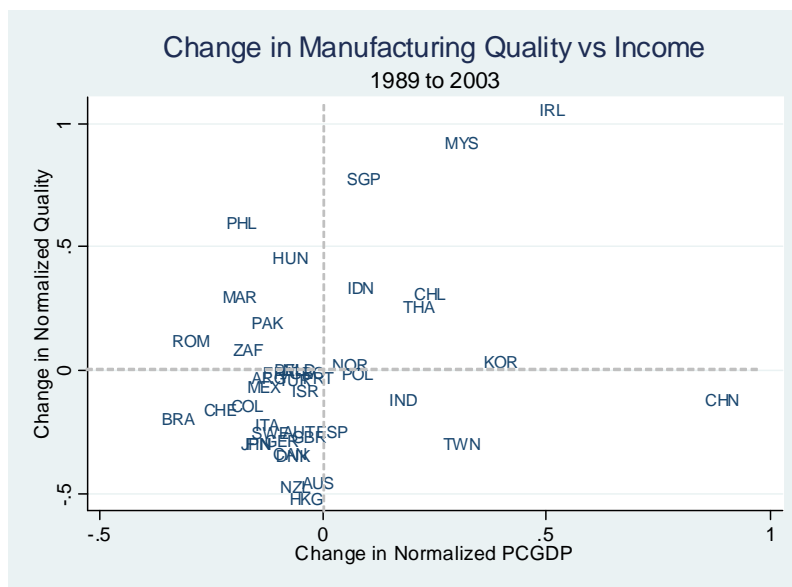


Figure 7: Change in Normalized Quality from 1989 to 2003 vs Change in Normalized Income 1989 to 2003

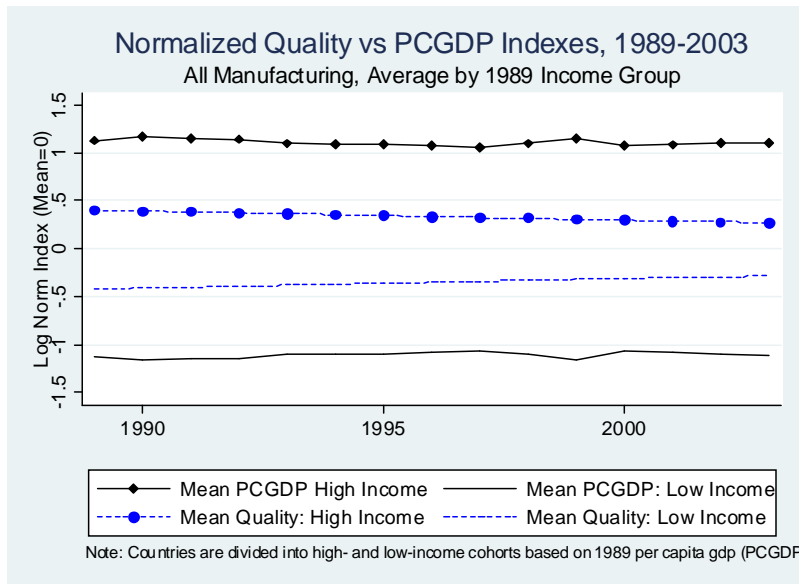


Figure 8: Evolution of Mean Normalized Quality and PCGDP for Countries with High and Low Income in 1989

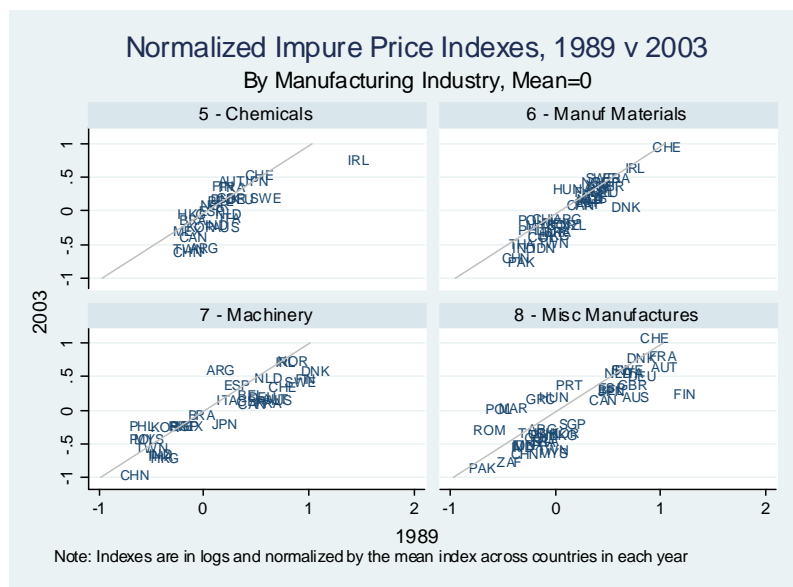


Figure 9: Normalized Log Impure Price Indexes Across Manufacturing Industries, 2003 vs 1989

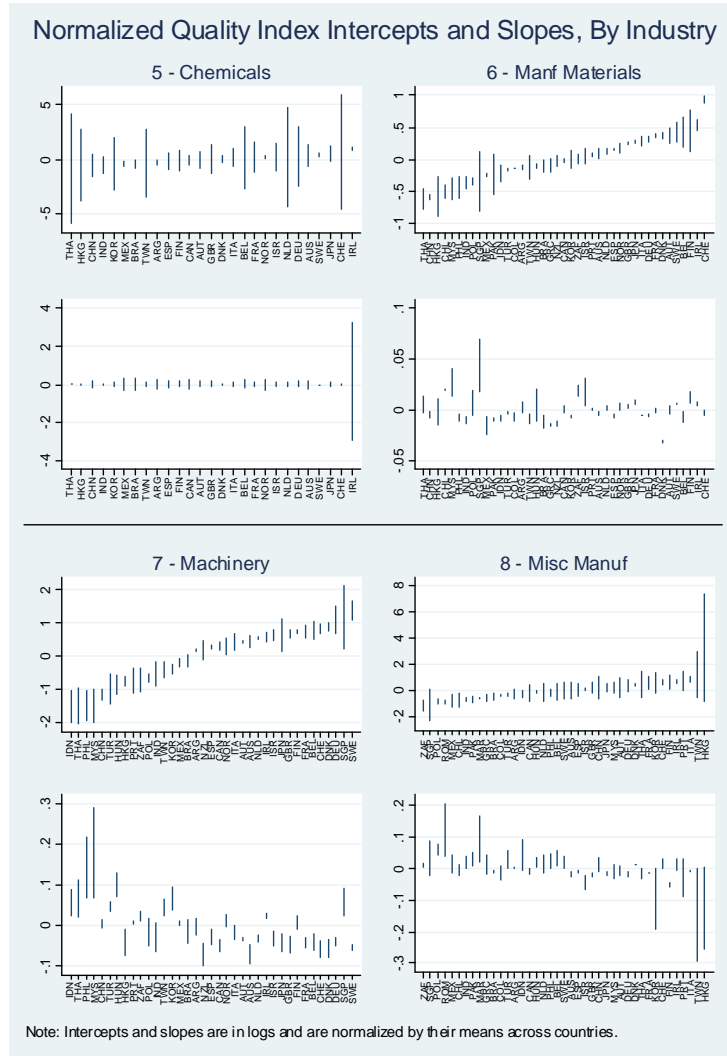


Figure 10: Normalized Log Quality Index Intercepts and Slopes, by Country and Manufacturing Industry

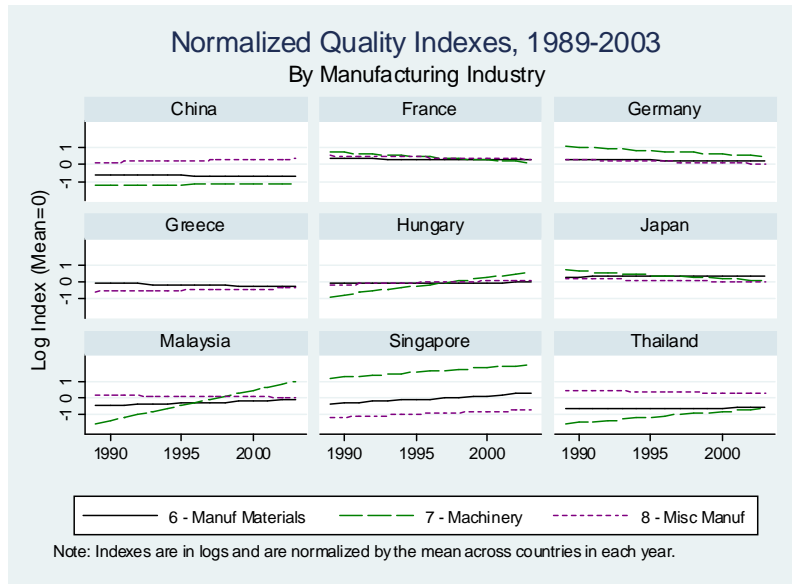


Figure 11: Normalized Log Quality Indexes Across Manufacturing Industries, 1989 to 2003

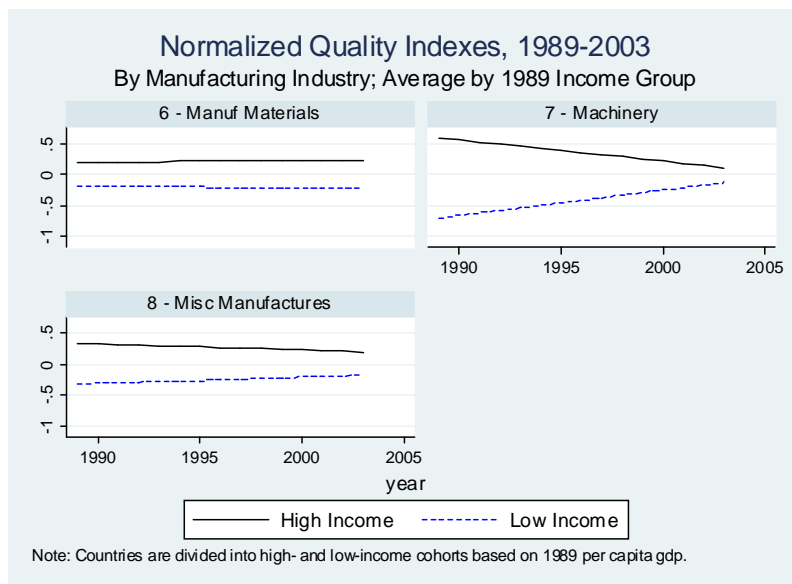


Figure 12: Evolution of Mean Normalized Quality for Countries with High and Low Income in 1989, by Industry