ESTIMATING CROSS-COUNTRY DIFFERENCES IN PRODUCT QUALITY*

Juan Carlos Hallak and Peter K. Schott

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We develop a method for decomposing countries' observed export prices into quality versus quality-adjusted components using information contained in their trade balances. Holding observed export prices constant, countries with trade surpluses are inferred to offer higher quality than countries running trade deficits. Our method accounts for variation in trade balances induced by both horizontal and vertical differentiation, and we use it to estimate the evolution of manufacturing quality for the world's top exporters from 1989 to 2003. We find that observed unit value ratios can be a poor approximation for relative quality differences, that countries' quality is converging more rapidly than their income, and that countries appear to vary in terms of displaying "high-quality" versus "low-price" growth strategies.

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I. Introduction

Theoretical and empirical research increasingly point to the importance of product quality in international trade and economic development.[1] 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 product 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).[2] 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 exporters can survive in the U.S. market even in the face of cost disadvantages.

Our method for identifying countries' product quality involves decomposing

1. 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 2008), countries' skill premia (Verhoogen 2008), 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).

2. 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).

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 higher product quality.[3]

A major contribution of the paper is to generalize this intuition to a setting where countries also are allowed to differ in the number of *unobserved* horizontal varieties they export in each product category (e.g., red versus blue men's wool sweaters). Horizontal differentiation is a standard aspect of recent trade models, and allowing for it helps explain why many products are exported by a wide range of countries. Incorporating it here 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

3. 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".

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.[4]

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 comparatively coarse industries relative to the much more disaggregated products (e.g., men's wool sweaters) at which some countries' export prices can be observed. To deal with this constraint we derive a theoretically appropriate price index that aggregates countries' observed product-level export prices up to the industry level. We refer to this index as the "Impure Price Index" because it is based on prices that are "contaminated" by quality. Our index has the useful property of being separable into quality versus quality-adjusted-price components, but it is developed under the potentially strong assumption that countries' quality is constant across products within industries. Thus, we are faced with an "aggregation trade-off": while product quality is more likely to be constant across products the more disaggregated the industry, data on countries' global net trade becomes more scarce as well as more susceptible to measurement error. Use of disaggregated industries may also be problematic if countries' use of intermediate inputs straddles the industry at which quality is being estimated; in this case, reported net trade in that industry fails to account for all of its inputs. In a pilot examination of this issue in the Data Appendix below, we find that apparel quality can be over-estimated for

4. Feenstra (1994) outlines a method for computing import price indexes that accounts for the introduction of new product varieties. (See also Broda and Weinstein 2004). Given its focus on changes in prices over time, that method 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.

countries that import textiles to produce apparel.

Even though the Impure Price Index comparing two countries' export prices is unobservable, we show that it is bounded by observable Paasche and Laspeyres indexes defined over their common exports to a third country (i.e., the United States). This result anchors a two-stage strategy for inferring countries' product 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 an 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 quality for overall manufacturing increases most dramatically for Ireland, Malaysia and Singapore over the sample period, and falls most dramatically for Hong Kong and New Zealand. Among countries that begin the sample period in the top tercile of quality, Australia and Japan experience the largest relative declines. We also show that our estimates of product quality and their evolution over time can deviate substantially from estimates of quality based on raw export prices. Indeed, changes in estimated relative quality 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. We also find greater narrowing in estimated quality differences than per capita GDP differences over our sample period. An interesting question for further research is the extent to which this quality convergence reveals a catching up in terms of technological knowledge by developing countries versus greater use of high-quality intermediate inputs from developed economies.

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.[5] 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.[6]

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.[7] Our approach is also different from an earlier

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

6. 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).

7. More recently, Khandelwal (2008) has developed a method for estimating quality based on the assumption of a nested logit demand system.

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 that literature, import quality increases when the composition of imports shifts toward high-quality product categories. Here, we take a within- rather than across-product view of quality variation.

Our results also relate well to recent efforts by Rodrik (2006), Hausmann, Hwang and Rodrik (2007) and others to estimate the extent to which the export quality of developing countries like China is equal to that of the world's most developed economies. Like Schott (2008) and Xu (2007), we find Chinese quality to be relatively low compared to developed countries across all years of our sample.

The paper is structured as follows. Section II outlines our assumptions about consumer demand and introduces the Impure and Pure Price indexes that will be the focus of our analysis. Section III shows that the unobservable Impure Price Index is bounded by observable Paasche and Laspeyres indexes. Section IV derives the relationship between the Pure Price Index and countries' sectoral net trade. Sections VI through VII describe the application of our method to identifying export quality trends for 43 large trading countries over the period 1989 to 2003. Section VIII concludes. Two appendixes attached to this paper provide proofs of our main propositions and an examination of quality by manufacturing industry. A web-based technical appendix contains estimation details and additional results.

II. 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.

II.A. Preferences

Varieties of goods are classified into product categories ("products" for short), which are in turn classified into sectors. Sectors are indexed by subscript $s = 1, ..., S$, while products (within sector s) are indexed by subscript $z = 1, ..., Z_s$. Product categories are the level of aggregation at which prices are observed while sectors are the level of aggregation at which countries' trade balances are observed and hence quality is estimated. In our empirical investigation below, products correspond to ten-digit U.S. Harmonized System (HS) categories while sectors are defined alternatively as All Manufacturing, one-digit SITC manufacturing industries or select two-digit SITC manufacturing industries. The theoretical framework presented here focuses on sector s.

There are K countries, indexed by superscript k. Preferences are represented by a two-tier utility function that incorporates consumer love of variety.[8] The upper tier is Cobb-Douglas while the lower tier is CES,

$$(1) \qquad U = \prod_{s=1}^{S} u_s^{b_s}, \qquad u_s = \left[\sum_{k=1}^{K} \sum_{z=1}^{Z_s} \left(\xi_z \lambda_s^k x_z^k \right)^{\frac{\sigma_s - 1}{\sigma_s}} n_z^k \right]^{\frac{\sigma_s}{\sigma_s - 1}}, \qquad \sigma_s > 1,$$

where n_z^k is the number of horizontally differentiated varieties of product z produced by country k, x_z^k is the quantity consumed per variety, and σ_s is the elasticity of substitution between varieties. For compactness, we omit subindexing z by s in the second summation of equation (1) and throughout the paper. We note that by indexing products instead of varieties, we implicitly assume symmetry across varieties of the same product.

The utility function includes two shifters, ξ_z and λ_s^k. The first shifter, ξ_z, varies across products but is constant across countries for a particular product. It captures consumers' valuation of the essential characteristics common to the

8. Homothetic preferences, although standard in the international trade literature, are potentially strong in this context as countries' demand for quality may vary with income.

heterogeneous varieties of a product. Consumers, for example, might have a higher preference for varieties of tables than chairs. The second shifter, λ_s^k, varies across countries and sectors, but is constant across products within a particular country and sector. It represents "quality", which we define as any attribute of a good (other than price and those already captured by ξ_z) for which all consumers are willing to pay more, and includes tangibles (e.g., durability) as well as intangibles (e.g., product image due to advertising). These assumptions, implicit in (1), are formalized as:

Assumption 1: $\xi_z^k = \xi_z, \qquad \forall k = 1, ..., K.$

Assumption 2: $\lambda_z^k = \lambda_s^k, \qquad \forall z = 1, ..., Z_s.$

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

II.B. Price and Quality Indexes

In this section we introduce the price and quality indexes that are the focus of our analysis. First define an aggregator of observed product prices produced in country k and sector s as[9]

$$P_s^k \equiv \left[\sum_z \overline{n}_z \xi_z^{\sigma_s - 1} \left(p_z^k\right)^{1-\sigma_s}\right]^{\frac{1}{1-\sigma_s}}, \qquad \overline{n}_z = \frac{1}{K}\sum_k \frac{n_z^k}{\frac{1}{Z_s}\sum_z n_z^k}.$$

We can then define the Impure Price Index (IPI) between countries k and k' as

$$P_s^{kk'} = P_s^k / P_s^{k'}. \tag{2}$$

P_s^k is a weighted average of country k's observed prices across products z in sector s, where each z is weighted according to the "world average" number of varieties ($\overline{n}_z$) and the demand shifter ($\xi_z^{\sigma_s-1}$) for that product.

The Impure Price Index is a summary measure of price variation between goods produced by countries k and k' in sector s. It has three features worth noting. First, because it is defined over observed prices it is "impure" in the sense that its prices are "contaminated" by quality. Second, it is transitive: choosing an arbitrary country, o, as numeraire, $P_s^{kk'}$ can always be recovered from the ratio $P_s^{ko}/P_s^{k'o}$. Finally, though unobservable due to its inclusion of unobserved variables such as the number of varieties countries export, this index can be estimated. In the next section, we show that the unobservable IPI is bounded by observable price indexes while in Section 5.1 we show how those bounds can be used to estimate the IPI. An alternate index based on n_z^k (rather than on $\overline{n}_z$) would have the advantage of being a subaggregate of the exact consumer price index and a more accurate predictor of countries' net trade. However, as will become clear later, the fact that this alternate index does not use weights that are common to all countries implies that it cannot be

9. To simplify notation, unless otherwise noted the subindexes under the summation sign range over all elements of the relevant set, e.g., $z = 1, ..., Z_s$ and $k = 1, ...K$.

bounded by observable price indexes.

We define a Quality Index, $\lambda_s^{kk'} = \lambda_s^k / \lambda_s^{k'}$, as the ratio of two countries' quality levels in sector s, and define a Pure Price Index (PPI), $\widetilde{P}_s^{kk'} = \widetilde{P}_s^k / \widetilde{P}_s^{k'}$, as the ratio of pure price aggregators, $\widetilde{P}_s^k \equiv \left[\sum_z \overline{n}_z \left(\widetilde{p}_z^k \right)^{1-\sigma_s} \right]^{\frac{1}{1-\sigma_s}}$. The Impure Price Index can be decomposed into the Quality Index and the Pure Price Index:

$$P_s^{kk'} = \lambda_s^{kk'} \widetilde{P}_s^{kk'}. \tag{3}$$

Estimating λ_s^{ko} is our main objective. Although both λ_s^{ko} and $\widetilde{P}_s^{ko}$ are unobservable, we show in Section 5.2 how they can be identified from estimates of P_s^{ko} and information on countries' net trade with the world in sector s.

III. Bounding the "Impure" Price Index

In this section we show that the unobserved Impure Price Index introduced above is bounded by observable Paasche and Laspeyres indexes defined over the prices (unit values) of country pairs' exports to a third country. This result is the basis of the strategy for estimating the IPI as outlined in Section 5.1. Our bounding of the Impure Price Index proceeds in two steps. First, we show that observable Paasche and Laspeyres indexes bound unobserved "cost-of-utility" indexes. Second, we show that these unobserved cost-of-utility indexes bound the unobserved Impure Price Index.

III.A. Paasche and Laspeyres Bounds on Cost-of-Utility Indexes

We define unobserved cost-of-utility indexes and use revealed preference to show that they are bounded by observed Paasche and Laspeyres indexes. Though the bounding of cost-of-utility indexes by Paasche and Laspeyres indexes is standard in the index number literature, our setup involves two complications. First, rather than concentrating on expenditures over the universe of goods in two different time periods, we focus on contemporaneous expenditures over subsets of the universe of goods purchased from a pair of exporting countries. Second, because we allow for horizontal differentiation, our cost-of-utility indexes need to deal with the number of varieties countries export – which need not be the same in the two countries.

We focus on countries' exports to a single "common importer", which we refer to as the United States given the data used in our empirical implementation. We note that the analysis would be identical were it to be applied to any other common importer, or to a set of importers. For ease of exposition, we assume in this section that all countries are "active" in (i.e., export to the United States) the same set of products, deferring discussion of the more general case of imperfect overlap to the Theory Appendix at the end of this paper. We summarize the implications of imperfect overlap for Proposition 1 after introducing the proposition below, and discuss the potential impact of imperfect overlap on our empirical analysis in Section 5.

Define vectors p_s^k and q_s^k to include, respectively, U.S. import prices and quantities for all products in sector s coming from country k. Stack these vectors across countries to form $\mathbf{p}_s$ and $\mathbf{q}_s$. Stack the latter vectors across sectors to form $\mathbf{p}$ and $\mathbf{q}$. Analogously, define vectors $\mathbf{n}$, $\boldsymbol{\lambda}$, and $\boldsymbol{\xi}$. A vector of per-variety consumption $\mathbf{x}$ is implicitly defined by $\mathbf{q}$ and $\mathbf{n}$. Finally, define $\mathbf{q}_s^{-k}$ as the complement of $\mathbf{q}_s^k$ with respect to $\mathbf{q}$. Vector $\mathbf{q}_s^{-k}$ includes import quantities in sector s from all countries other than k, and also import quantities in all other sectors from all countries (including k).

For country k of country pair kk', we define the constrained expenditure (or import) function $m_{s,k}(\mathbf{p}_s^{k''}, \mathbf{q}_s^{-k}, \mathbf{n}, \boldsymbol{\lambda}, \boldsymbol{\xi}, u)$ as the solution to the problem

$$(4) \qquad \min_{\widehat{q}_s^k} \ \mathbf{p}_s^{k''}\widehat{\mathbf{q}}_s^k \qquad s.t. \quad U(\widehat{\mathbf{q}}_s^k, \mathbf{q}_s^{-k}, \mathbf{n}, \boldsymbol{\lambda}, \boldsymbol{\xi}) = u, \qquad k'' = 1, ..., K$$

where U is the representative consumer utility function.[10] This function represents the minimum expenditure on varieties in sector s imported from country k that the consumer would be required to make in order to attain utility level u if import prices of those varieties were $\mathbf{p}_s^{k''}$ (rather than $\mathbf{p}_s^k$), holding constant the actual values of $\mathbf{q}_s^{-k}, \mathbf{n}, \boldsymbol{\lambda}, \boldsymbol{\xi}$.

To obtain an explicit functional form for $m_{s,k}$, we use the preferences outlined in equation (1). Define $u_s^k \equiv \left[\sum_z n_z^k \left(\xi_z \lambda_s^k x_z^k\right)^{\frac{\sigma_s - 1}{\sigma_s}}\right]^{\frac{\sigma_s}{\sigma_s - 1}}$ only over varieties exported by country k in sector s. The separability of the utility function in (1) implies that U can be written as a function of u_s^k and a function of arguments held constant in problem (4), $\overline{u}_s^{-k}$. Since U is strictly increasing in u_s^k, there is a single value of this variable, $\widehat{u}_s^k$, such that $U\left(\widehat{u}_s^k, \overline{u}_s^{-k}\right) = u$. Then, problem (4) reduces to choosing the per-variety quantities x_z^k that minimize $\sum_z n_z^k p_z^{k''} x_z^k$ subject to $u_s^k = \widehat{u}_s^k$. The solution to this problem is the product of a CES aggregator measuring the unit cost of utility and the target level of utility, $\widehat{u}_s^k$[11]

$$(5) \qquad m_{s,k}(\mathbf{p}_s^{k''}, \mathbf{q}_s^{-k}, \boldsymbol{\lambda}, \boldsymbol{\xi}, u) = \left[\sum_z n_z^k \left(\widetilde{p}_z^{k''} \frac{\lambda_s^{k''}}{\lambda_s^k}\right)^{1-\sigma_s}\right]^{\frac{1}{1-\sigma_s}} \widehat{u}_s^k.$$

By revealed preference, $m_{s,k}(\mathbf{p}_s^k, \mathbf{q}_s^{-k}, \mathbf{n}, \boldsymbol{\lambda}, \boldsymbol{\xi}, u) = \mathbf{p}_s^k\mathbf{q}_s^k$. However, if prices were $\mathbf{p}_s^{k'}$ instead of $\mathbf{p}_s^k$, the minimum import expenditure would be equal to or lower than $\mathbf{p}_s^{k'}\mathbf{q}_s^k$, because the amount $\mathbf{p}_s^{k'}\mathbf{q}_s^k$ is sufficient to attain utility u but $\mathbf{q}_s^k$ is not necessarily optimal given $\mathbf{p}_s^{k'}$. Hence, $m_{s,k}(\mathbf{p}_s^{k'}, \mathbf{q}_s^{-k}, \mathbf{n}, \boldsymbol{\lambda}, \boldsymbol{\xi}, u) \leq$

10. Neary and Roberts (1980) and Anderson and Neary (1992) use the constrained expenditure function to analyze consumption choices under rationing.

11. It is here where Assumptions 1 and 2 are critical. In equation (5) we use these assumptions to derive $\frac{p_z^{k''}}{\lambda_z^k \xi_z^k} = \frac{p_z^{k''}}{\lambda_z^{k''} \xi_z^{k''}} \frac{\lambda_z^{k''} \xi_z^{k''}}{\lambda_z^k \xi_z^k} = \widetilde{p}_z^{k''} \frac{\lambda_s^{k''}}{\lambda_s^k}$.

$\mathbf{p}_s^{k'}\mathbf{q}_s^k$. Using these results, we obtain

$$M_{s,k}^{kk'} \equiv \frac{m_{s,k}(\mathbf{p}_s^k, \mathbf{q}_s^{-k}, \mathbf{n}, \boldsymbol{\lambda}, \boldsymbol{\xi}, u)}{m_{s,k}(\mathbf{p}_s^{k'}, \mathbf{q}_s^{-k}, \mathbf{n}, \boldsymbol{\lambda}, \boldsymbol{\xi}, u)} \geq \frac{\mathbf{p}_s^k \mathbf{q}_s^k}{\mathbf{p}_s^{k'} \mathbf{q}_s^k} = H_s^{kk'}. \tag{6}$$

Inequality (6) displays a standard result in index number theory stating that the cost-of-utility price index $M_{s,k}^{kk'}$ is larger than a Paasche price index, $H_s^{kk'}$, defined over the observed prices of the country pair's exports to the U.S. in sector s. We note that the Paasche index is defined here in a cross-sectional rather than a time-series context. $M_{s,k}^{kk'}$ captures the change in minimum expenditure on country k's varieties (in sector s) that would be necessary to maintain utility u if import prices of those varieties changed from $\mathbf{p}_s^{k'}$ to $\mathbf{p}_s^k$, holding constant their number and characteristics (including quality), and the number, characteristics and quantity consumed of all other goods.

We can combine equation (5) with inequality (6) to obtain

$$\ln H_s^{kk'} \leq \ln M_{s,k}^{kk'} = \ln P_s^{kk'} + \ln \phi_{s,k}^{kk'}, \qquad \phi_{s,k}^{kk'} \equiv \left[\frac{\sum_z n_z^k \left(\frac{\widetilde{p}_z^k}{\widetilde{P}_s^k} \right)^{1-\sigma_s}}{\sum_z n_z^k \left(\frac{\widetilde{p}_z^{k'}}{\widetilde{P}_s^{k'}} \right)^{1-\sigma_s}} \right]^{\frac{1}{1-\sigma_s}}. \tag{7}$$

In a similar manner, we can focus alternatively on imports from country k' to obtain

$$M_{s,k'}^{kk'} \equiv \frac{m_{s,k'}(\mathbf{p}_s^k, \mathbf{q}_s^{-k'}, \mathbf{n}, \boldsymbol{\lambda}, \boldsymbol{\xi}, U)}{m_{s,k'}(\mathbf{p}_s^{k'}, \mathbf{q}_s^{-k'}, \mathbf{n}, \boldsymbol{\lambda}, \boldsymbol{\xi}, U)} \leq \frac{\mathbf{p}_s^k \mathbf{q}_s^{k'}}{\mathbf{p}_s^{k'} \mathbf{q}_s^{k'}} = L_s^{kk'}, \tag{8}$$

where $L_s^{kk'}$ is a Laspeyres price index. This is another standard result, which states that the cost-of-utility index $M_{s,k'}^{kk'}$ is bounded from above by a Laspeyres price index. Using the explicit functional form for $m_{s,k'}$, we obtain

(9)

$$\ln L_s^{kk'} \geq \ln M_{s,k'}^{kk'} = \ln P_s^{kk'} + \ln \phi_{s,k'}^{kk'}, \qquad \phi_{s,k'}^{kk'} \equiv \left[\frac{\sum_z n_z^{k'} \left(\frac{\widetilde{p}_z^k}{\widetilde{P}_s^k} \right)^{1-\sigma_s}}{\sum_z n_z^{k'} \left(\frac{\widetilde{p}_z^{k'}}{\widetilde{P}_s^{k'}} \right)^{1-\sigma_s}} \right]^{\frac{1}{1-\sigma_s}}.$$

Equations (7) and (9) 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 one side.[12] Although a standard result in the index number literature shows that the cost-of-utility index for a consumer with homothetic preferences is independent of the utility level – and bounded both above *and* below – our allowance for horizontal differentiation yields two cost-of-utility indexes because $M_{s,k}^{kk'}$ and $M_{s,k'}^{kk'}$ are defined over different numbers of varieties, i.e., n_z^k and $n_z^{k'}$, respectively. $M_{s,k}^{kk'}$ and $M_{s,k'}^{kk'}$ would be equal if, for example, the number of varieties in countries k and k' were proportional to one another for every product category.[13]

III.B. Paasche and Laspeyres Bounds on the Impure Price Index

To bound the (unobservable) Impure Price Index by the observable Paasche and Laspeyres indexes via the cost-of-utility indexes defined above, we must

12. 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), inequalities (7) and (9) also hold if $M_{s,c}^{kk'}, M_{s,d}^{kk'}, H_s^{kk'}, L_s^{kk'}$ are alternatively defined using free-on-board (FOB) prices – i.e., exclusive of customs, insurance and freight charges – as all terms are simply scaled by the relative trade costs between countries k and k' and the United States. As noted in Section 5, we use *fob* import unit values to measure U.S. trading partners' export prices in our empirical analysis.

13. Note also that the indexes $H_s^{kk'}, L_s^{kk'}, M_{s,k}^{kk'}$ and $M_{s,k'}^{kk'}$ all weight prices in the numerator and in the denominator with the same weights, respectively $\mathbf{q}_s^k, \mathbf{q}_s^{k'}, n_z^k$, and $n_z^{k'}$. Our ability to bound $P_s^{kk'}$ with those indexes in the next section depends crucially on $P_s^{kk'}$ also having weights, $\overline{n}_z$, that are common in the numerator and denominator.

show that $\ln \phi_{s,k}^{kk'} \leq 0$ and $\ln \phi_{s,k'}^{kk'} \geq 0$ so that $H_s^{kk'} \leq M_{s,k}^{kk'} \leq P_s^{kk'} \leq M_{s,k'}^{kk'} \leq L_s^{kk'}$. In this section, we outline assumptions that are sufficient for these conditions to hold.

Our first step is to decompose the number of varieties countries produce into three meaningful parts. Let $\overline{n}_s^k = \frac{1}{Z_s} \sum\limits_z n_z^k$ be country k's average number of varieties across product categories in sector s. Let $\overline{n}_z^{kk'} = \frac{1}{2}\left(\frac{n_z^k}{\overline{n}_s^k} + \frac{n_z^{k'}}{\overline{n}_s^{k'}}\right)$ be the (normalized) average number of varieties of product z in sector s across members of the country pair. Then, the (normalized) number of varieties a country produces can be expressed as the sum of three terms:

$$\frac{n_z^k}{\overline{n}_s^k} = \overline{n}_z + \widetilde{n}_z^{kk'} + \widetilde{n}_z^{k,kk'}. \tag{10}$$

The first term is the world average for product z introduced in Section 2.[14] The second term is the "country-pair excess variety" in product z relative to the world average, $\widetilde{n}_z^{kk'} = \overline{n}_z^{kk'} - \overline{n}_z$, which captures the extent to which the average number of varieties in country pair kk' is above or below the world average. The third term is country k's "bilateral excess variety" for product z relative to kk''s average, $\widetilde{n}_z^{k,kk'} = \frac{n_z^k}{\overline{n}_s^k} - \overline{n}_z^{kk'}$. We note that $\sum\limits_z \widetilde{n}_z^{kk'} = 0$, $\sum\limits_z \widetilde{n}_z^{k,kk'} = 0$ and $\sum\limits_z \widetilde{n}_z^{k',kk'} = 0$: that is, the pair kk' cannot have positive country-pair excess variety in all z and neither country can have positive bilateral excess variety in all z. Finally, $\sum\limits_{k''=k,k'} \widetilde{n}_z^{k'',kk'} = 0$: k and k' cannot both have positive bilateral excess variety in the same z.

Our second step is to define the (normalized) bilateral difference in countries'

14. On notation: recall that implicit in our use of the index z is the understanding that it pertains to the z within sector s. Thus, all terms in equation (10) refer to a particular sector s.

pure prices in product z as

$$(11) \qquad \Delta\widetilde{p}_z^{kk'} = \left(\frac{\widetilde{p}_z^k}{\widetilde{P}_s^k}\right)^{1-\sigma_s} - \left(\frac{\widetilde{p}_z^{k'}}{\widetilde{P}_s^{k'}}\right)^{1-\sigma_s} , \qquad (\Delta\widetilde{p}_z^{k'k} = -\Delta\widetilde{p}_z^{kk'}) \quad .$$

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

Assumption 3 states that country k relative to country k' will tend to have positive bilateral excess variety in those products in which it has a lower relative pure price (the operator cov_s denotes sample covariances defined over all z in sector s).

Assumption 3: $cov_s\left(\widetilde{n}_z^{k,kk'}, \Delta\widetilde{p}_z^{kk'}\right) = cov_s\left(\widetilde{n}_z^{k',kk'}, \Delta\widetilde{p}_z^{k'k}\right) \geq 0$

This assumption is motivated by 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 products 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 is a negative function of relative pure (or quality-adjusted) prices.[15]

15. In a multi-country set up, the relative number of varieties between two countries is also determined by the pure prices of third countries. Therefore, Assumption 3 implicitly imposes

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

Assumption 4: $cov_s\left(\widetilde{n}_z^{kk'}, \Delta\widetilde{p}_z^{kk'}\right) = 0$

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.

With assumptions 3 and 4 as well as our earlier assumptions about consumer utility, we obtain the main result of this section:

PROPOSITION 1. Under Assumptions 1 through 4, for any two countries k and k', the (unobservable) Impure Price Index is bounded by the (observable) Paasche and Laspeyres indexes:

$$\ln H_s^{kk'} \leq \ln P_s^{kk'} \leq \ln L_s^{kk'}$$

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. As noted above, it assumes all countries are active in the same set of products. As discussed in the Theory Appendix at the end of this paper, the more general case of imperfect overlap may result in violations of Proposition 1. We show however that such violations are less likely when the number of mismatched products is low and when mismatched products are more evenly distributed across countries in a pair. As discussed further in Section 5, we attempt to mitigate the possibility of such violations in our empirical analysis by excluding country pairs with few export

that bilateral price effects dominate over price effects with respect to third countries. We thank a referee for making this point.

products in common and by considering subsets of our sample countries which overlap in greater numbers of products.

IV. 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 k to country k' requires paying iceberg trade costs of $\tau_s^{kk'}$. Therefore, $p_z^k \tau_s^{kk'}$ is the import price of product z in country k'. Given the CES preferences over products in sector s outlined above, it is easy to derive country k's bilateral export and import flows in sector s with every other country. Summing export flows over all partners $k' \neq k$, we obtain the value of country k's exports,

$$Exports_s^k = \sum_{k' \neq k} \left[\sum_z \frac{n_z^k \left(\widetilde{p}_z^k \tau_s^{kk'} \right)^{1-\sigma_s}}{\left(G_s^{k'}\right)^{1-\sigma_s}} \right] b_s E^{k'} \tag{12}$$

where $G_s^{k'}$ is a consumption-based price aggregator and $\left(G_s^{k'}\right)^{1-\sigma_s} = \sum_{k''} \sum_z n_z^{k''} \left(\widetilde{p}_z^{k''} \tau_s^{k''k'} \right)^{1-\sigma_s}$. $E^{k'}$ is the expenditure of country k' and equals its income ($Y^{k\prime}$) minus its trade balance ($T^{k\prime}$). The expression in brackets in equation (12) is country k's share in country $k\prime$'s sectoral expenditure. Prices and quality levels affect this share only through their ratio, $\widetilde{p}_z^k$.[16]

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

$$Imports_s^k = \left[1 - \sum_z \frac{n_z^k \left(\widetilde{p}_z^k \right)^{1-\sigma_s}}{\left(G_s^k\right)^{1-\sigma_s}} \right] b_s E^k. \tag{13}$$

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

Subtracting equation (13) from equation (12), we obtain country k's net trade with the world in sector s, T_s^k, as a proportion of its expenditure in the sector,

$$\frac{T_s^k}{b_s E^k} = \left(\sum_z \frac{n_z^k}{E^k} \left(\widetilde{p}_z^k\right)^{1-\sigma_s} \right) \exp\left(\tau_s^k\right) - 1 \tag{14}$$

where $\tau_s^k = \ln\left(\sum_{k'} E^{k'} \left(\frac{\tau_s^{kk'}}{G_s^{k'}} \right)^{1-\sigma_s} \right)$.

The summary measure of trade costs, τ_s^k, captures bilateral trade costs between all country pairs. First, it includes all outbound bilateral trade costs for country k. Those costs, $\tau_s^{kk'}$, enter directly, so that τ_s^k is smaller the higher are those costs. Second, via G_s^k, τ_s^k also includes all inbound bilateral trade costs for country k, $\tau_s^{k'k}$, so that τ_s^k is larger the larger are those costs. Finally, all other bilateral trade costs enter indirectly through countries' consumption price indexes, $G_s^{k'}$, dampening the negative effect of outbound bilateral trade costs. As a result, net trade of country k is higher the higher are trade costs between third countries.[17]

Equation (14) shows that a country's net trade (per expenditure in the sector) is a function of its pure prices and numbers of varieties, its total expenditure, and a summary measure of its bilateral trade costs, τ_s^k. Our objective is to derive a version of equation (14) that reduces the dimensionality of unobservables and that can be related to the estimated Impure Price Index.

To achieve this objective, define country k's "multilateral excess variety" in product z as $\widetilde{n}_z^k = \frac{n_z^k}{\overline{n}_s^k} - \overline{n}_z$, where $\sum_z \widetilde{n}_z^k = 0, \forall k = 1, ..., K$. The covariance between multilateral excess variety and (normalized) pure prices can be expressed as the sum of a common component across countries (φ_s) and a mean-zero,

17. See Anderson and van Wincoop (2003) for a detailed discussion of the effects of trade costs on trade flows in a related setting.

country-specific idiosyncratic component[18]

$$\text{(15)} \qquad cov_s\left[\widetilde{n}_z^k, \left(\widetilde{p}_z^k/\widetilde{P}_s^k\right)^{1-\sigma_s}\right] = \varphi_s + \mu_s^k,$$

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 products within sectors.[19]

Assumption 5: $\overline{n}_s^k/Y^k = \left(\widetilde{P}_s^k\right)^{-\eta_s}, \quad \forall k = 1, ..., K, \qquad \eta_s \geq 0.$

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.[20]

The following Proposition describes the main result of this section.

Proposition 2. Under Assumption 5, country k's net trade in sector s (above the sector's proportional share in total net trade) can be approximated as a log-linear function of $\widetilde{P}_s^k$

$$\text{(16)} \qquad \frac{T_s^k - b_s T^k}{E^k} = \Upsilon_s + \gamma_s \ln \widetilde{P}_s^k + b_s \tau_s^k + \iota_s^k$$

18. Note that this characterization does not impose any restriction on the covariance. For estimation, we will assume that μ_s^k and the instrumental variable are uncorrelated.

19. In fact, in a coarse check of this assumption discussed further in the web-based technical appendix, we find a negative relationship between our estimated pure price indexes and the number of ten-digit HS products within manufacturing sectors that countries export to the United States (normalized by GDP) over our sample period.

20. This relationship abstracts from home market effects or "multilateral" effects such as being close to low- or high-pure price countries, which could affect the number of varieties that countries produce.

where

$$\Upsilon_s = b_s Z_s \varphi_s, \quad \gamma_s = b_s \left(1 - \sigma_s - \eta_s\right) < 0, \quad \iota_s^k = b_s Z_s \mu_s^k,$$

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 – above the sector's share in total net trade – should be larger the lower are its pure prices.

In addition to pure prices, trade costs also influence net trade. Proposition 2 characterizes this influence. Since the proposition captures the impact of trade costs on net trade *conditional on pure prices*, it does 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 (16). 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 (16) can be interpreted as a demand function, where the sectoral net trade with the world is the "quantity" variable, $\widetilde{P}_s^k$ is the "price" variable, and τ_s^k is a demand shifter. The first term captures movements along the demand curve: higher pure prices of country k 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.

We use countries' trade *balances* with the *world* as the "quantity" indicator in our method to mitigate our inability to control for unobserved components of bilateral trade costs, i.e., information costs, idiosyncratic transport costs, and non-tariff barriers associated with commercial policy. By using trade *balances* rather than either exports or imports alone, we cause unobserved components of countries' trade costs that affect both exports and imports in a country pair to cancel out. By using countries' trade balances with the *world*, i.e. summing a country's trade flows across all of its trading partners, we average out the impact of unobserved, idiosyncratic components of bilateral trade costs.[21,22] Still, unobserved components of trade costs that are neither canceled out by using trade balances nor averaged out by using trade balances with the world will inappropriately feed into our estimates of quality.

V. Estimation

In this section we demonstrate how our theoretical results can be used to estimate U.S. trading partners' relative manufacturing 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 dataset creation to a separate, web-based technical appendix.[23]

21. Khandelwal (2008), by contrast, relies on "demand" information contained in the imports of a single trading partner (the United States). An advantage of that approach is that U.S. imports can be observed at a more disaggregate level than world trade. A disadvantage is that, for the reasons noted above, one-way flows to a single country are likely to be substantially more sensitive to mismeasurement of trade costs than countries trade balances with the world.

22. We discuss results based on exports to the United States as an alternate measure of "demand" in the web-based technical appendix.

23. Datasets and computer code developed to generate our results are also available with this web-based technical appendix.

V.A. *Estimation of First-Stage Impure Price Indexes*

The first stage of the estimation uses Proposition 1 to estimate each country's Impure Price Index, $\widehat{P}_s^{ko}, \forall k \neq o$, where country o is the numeraire country (without loss of generality) and hats over variables denote estimates. For generic country pair k and k', the estimated indexes $\widehat{P}_s^{ko}$ and $\widehat{P}_s^{k'o}$ implicitly determine the bilateral index $\widehat{P}_s^{kk'} = \widehat{P}_s^{ko}/\widehat{P}_s^{k'o}$. This index should satisfy the Paasche and Laspeyres bounds for that country pair, as outlined in Proposition 1. Similarly, for K trading partners, the $K-1$ estimated Impure Price Indexes $\widehat{P}_s^{ko}$, $\forall k \neq o$, implicitly determine $K(K-1)$ bilateral indexes $\widehat{P}_s^{kk'}, \forall(k, k')$, 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 misrecorded 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 $H_s^{*kk'}$ and $L_s^{*kk'}$, respectively. We assume that the observed indexes, $H_s^{kk'}$ and $L_s^{kk'}$, depart from the true indexes by a multiplicative error: in logs, $\ln H_s^{kk'} = \ln H_s^{*kk'} + \varrho_{h,s}^{kk'}$ and $\ln L_s^{kk'} = \ln L_s^{*kk'} + \varrho_{l,s}^{kk'}$. We also assume that each error is distributed normally, with mean zero and standard deviation ψ_s, and that the errors for each bound are independent both of each other and of error terms for other bilateral pairs.[24]

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

$$\ln P_s^{kk'} \geq \ln H_s^{*kk'} \Rightarrow \varrho_{h,s}^{kk'} \geq \ln H_s^{kk'} - \ln P_s^{kk'} \tag{17}$$

$$\ln P_s^{kk'} \leq \ln L_s^{*kk'} \Rightarrow \varrho_{l,s}^{kk'} \leq \ln L_s^{kk'} - \ln P_s^{kk'}. \tag{18}$$

24. 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.

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 (17) and (18) are jointly satisfied for each country pair $\{k, k'\}$. This criterion implies maximizing the function

$$\ln L = \sum_k \sum_{kk'>k} \left\{ \ln \left[1 - \Phi \left(\frac{\ln H_s^{kk'} - \ln P_s^{kk'}}{\psi_s} \right) \right] + \ln \Phi \left(\frac{\ln L_s^{kk'} - \ln P_s^{kk'}}{\psi_s} \right) \right\}$$

where Φ is the cumulative normal.

Intuition for this estimator is provided in Figure I, which considers the Paasche-Laspeyres interval for a single country pair k and k', defined by $\ln H_s^{kk'}$ and $\ln L_s^{kk'}$. In the figure, two cumulative normal distributions, each with standard deviation ψ_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 $\ln \widehat{P}_s^{kk'} = \ln \widehat{P}_s^{ko} - \ln \widehat{P}_s^{k'o}$ along the horizontal axis in the figure. According to equation (17), the height of the cumulative normal distribution to the left of $\ln \widehat{P}_s^{kk'}$ indicates the likelihood that the true Paasche index is lower than the estimated bilateral index, that is, $\ln H_s^{*kk'} < \ln \widehat{P}_s^{kk'}$. Likewise, using equation (18), the height of the cumulative normal to the right of $\ln \widehat{P}_s^{kk'}$ indicates the likelihood that the true Laspeyres index is greater than the estimated bilateral index, that is, $\ln L_s^{*kk'} > \ln \widehat{P}_s^{kk'}$. Choosing a particular value for $\ln \widehat{P}_s^{kk'}$ 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 $\ln \widehat{P}_s^{kk'}$ is within the true bilateral Paasche and Laspeyres bounds, only taking into account the bounds of this particular country pair, then $\ln \widehat{P}_s^{kk'}$ would lie in the middle of the interval and be equivalent to the well-known Fisher index. However, because the choices of $\ln \widehat{P}_s^{ko}$ and $\ln \widehat{P}_s^{k'o}$, which determine $\ln \widehat{P}_s^{kk'}$ for this country pair, also influence the fit of all other country pairs in which either country k or k' 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 k and k'. For that reason, $\ln \widehat{P}_s^{kk'}$ is drawn off-center in the interval depicted in Figure I.

Our estimator has the advantage of penalizing 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).[25]

V.A..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 on U.S. import data given its level of detail and availability for such a long time horizon, but note that our method can be generalized to include data from other countries, which could be used to generate additional Paasche and Laspeyres bounds that

25. In the web-based technical appendix, we compare our estimator to three alternatives: a quadratic penalty function centered at the midpoint of each country pair's interval; a function that only penalizes estimates outside the interval; and an index proposed by Hummels and Klenow (2005) which compares countries' prices to those of the world over the set of goods they have in common with the world. We find that the first two alternatives yield IPI and quality estimates very similar to those reported below. Results using the third alternative vary more substantially from those reported below. However, the goodness of fit of that alternative, i.e., the percent of first-stage Impure Price Index estimates that lie within the Paasche-Laspeyres bounds, is considerably lower, thus supporting our choice of the estimator defined in the main text.

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.[26]

We compute the unit value, or "price", of export product z from source country k, p_z^k, by dividing free-on-board import value (v_z^k) by import quantity (q_z^k), $p_z^k = v_z^k/q_z^k$, where free-on-board refers to import values that are exclusive of customs, insurance and freight charges.[27] Examples of the units employed to classify products include dozens of men's 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 error (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 four additional criteria. First, a (more stringent) *Relevance Constraint* mandates that country-product-year observations must have quantity greater than 25 and value (in 1989 dollars) greater than $50,000. Second, a *Presence*

26. This assumption may not be innocuous. In principle, it could be tested by comparing the results of this section to results based on other countries' data

27. 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.

Constraint requires country-product observations to appear in more than two years of the sample. Third, a *Country-Pair Overlap Constraint* insists that, 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. Finally, a *Unit-Value Dispersion Constraint* requires that country-product-year observations be excluded if the country's adjusted[28] 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 impose a final constraint that data required for both the first and second stage cannot be missing for more than three years of the sample period. After all screens are implemented, we are left with 43 countries, which constitute the sample we use in the remainder of the paper.

The costs and benefits of screening the raw data can be discerned from Table I. 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 I 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 I show that secondary screening also reduces country and country-product participation in the sample, lowering

28. 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 implement this screen, 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.

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., $L_s^{kk'} < H_s^{kk'}$) in all three screens; for our preferred sample, just 0.6 percent of bounds are ordered incorrectly. We exclude those bounds from our 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 ($\ln L_s^{kk'} - \ln H_s^{kk'}$) 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 from 0.74) versus import value (to 97.8 from 88.8 percent).

The left-hand panel of Table II 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.

As highlighted in Section 3 and discussed further in the Theory Appendix below, the imperfect overlap of export products between countries induces potential violations to Proposition 1. Such violations might generate composition bias in the estimates of the Impure Price Indexes and, as a result, in estimates of the Quality Indexes. Further, growth in the product coverage of countries' exports might change the extent of bias over time, also affecting the estimated time trends. As noted above in Section 5.1.1 we attempt to mitigate the influence of composition bias via the use of the *Country-Pair Overlap* constraint when screening the raw data.[29] Below, we also compare the quality rankings of the thirty largest exporters in our sample to alternate estimates derived from restricting the analysis to just those thirty countries. Since these thirty countries exhibit substantially higher export-product overlap than all countries in the base sample, our finding of similar relative quality in both estimations suggests that composition bias, if present, is limited.

V.A..2 First-Stage Results

The right-hand panel of Table II 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, $\widehat{\psi}_s$, 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

29. Data restrictions prevent implementation of other potential solutions to this problem. We cannot restrict analysis to a set of continually exported country-products, for example, due to numerous changes to Harmonized System product classification codes over the sample period.

in 2003.

Estimation of the first stage yields an Impure Price Index for each country relative to the numeraire country. In Figure II, we report *normalized* log Impure Price Indexes for all countries for the first and last years of the sample. This normalization involves subtracting the mean log index across countries from every country's estimated log Impure Price Index, by year

$$\ln \widehat{P}_{st}^{k,Mean} = \ln \widehat{P}_{st}^{ko} - \frac{1}{K}\sum_{k'} \ln \widehat{P}_{st}^{k'o}. \tag{19}$$

In particular, the normalized Impure Price Index for the numeraire country (Switzerland), $\ln \widehat{P}_{st}^{o,Mean}$, is equal to $-\frac{1}{K}\sum_{k} \ln \widehat{P}_{st}^{k'o}$.

Estimated Impure Price Indexes generally accord with expectations. In the figure, countries nearer the lower left corner 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 (HUN) 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 IV.

V.B. Estimation of Second-Stage Quality Indexes

The second stage of our estimation uses Proposition 2 to recover information about countries' relative quality from their first-stage estimated Impure Price Indexes. First, we sum and subtract $\gamma_s \ln \widetilde{P}_s^o$ to the right-hand side of equation

(16) to express it as a function of the Pure Price Index (relative to numeraire o) rather than of the price aggregator $\widetilde{P}_s^k$. Then, since we calculate b_s from data, we take the trade cost term to the left-hand-side. Finally, we use $\ln \widetilde{P}_s^{ko} = \ln P_s^{ko} - \ln \lambda_s^{ko}$ to rewrite this equation as

$$\widetilde{T}_{st}^k = \Upsilon'_{st} + \gamma_s \ln \widehat{P}_{st}^{ko} - \gamma_s \ln \lambda_{st}^{ko} + \gamma_s \kappa_{st}^{ko} + b_s Z_s \mu_{st}^k \tag{20}$$

where $\widetilde{T}_{st}^k = \left(T_{st}^k - b_s T_t^k\right)/E_t^k - b_s \tau_{st}^k$, t indexes time periods, $\Upsilon'_{st} = \Upsilon_{st} + \gamma_s \ln \widetilde{P}_s^o$, and $\kappa_{st}^{ko} = \ln P_{st}^{ko} - \ln \widehat{P}_{st}^{ko}$ is the estimation error from the first stage. The last three terms in equation (20) are unobservable and create a compound error term that includes: countries' product quality relative to the numeraire country (λ_{st}^{ko}); the estimation error in the first stage (κ_{st}^{ko}); and the idiosyncratic component of the covariance between excess variety and pure prices (μ_{st}^k) from equation (15). Assuming that this compound error term is uncorrelated with the regressors is untenable. In particular, the quality component λ_{st}^{ko} 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,

$$\ln \lambda_{st}^{ko} = \alpha_{0s}^{ko} + \alpha_{1s}^{ko} t + \varepsilon_{st}^{ko} \tag{21}$$

where α_{0s}^{ko} and α_{1s}^{ko} are a country fixed effect and the slope of a country-specific time trend, respectively, and ε_{st}^{ko} 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

(20), we obtain our second-stage estimating equation

$$\widetilde{T}_{st}^{k} = \Upsilon'_{st} + \gamma_s \ln \widehat{P}_{st}^{ko} - \zeta_{0s}^{ko} - \zeta_{1s}^{ko} t + \upsilon_{st}^{ko} \tag{22}$$

where $\zeta_{0s}^{ko} = \gamma_s \alpha_{0s}^{ko}$ and $\zeta_{1s}^{ko} = \gamma_s \alpha_{1s}^{ko}$ are country fixed effects and time trends, respectively, and $\upsilon_{st}^{ko} = \gamma_s(\kappa_{st}^{ko} - \varepsilon_{st}^{ko}) + b_s Z_s \mu_{st}^{k}$ is the error term. Note that the term on the left-hand-side 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 (22) 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 ε_{st}^{ko} and $\widehat{P}_{st}^{ko}$ 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 $\ln \widehat{P}_{st}^{ko}$, *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 $\widehat{P}_{st}^{ko}$. 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 $\widetilde{P}_{st}^{ko}$. Since $\widetilde{P}_{st}^{ko}$ is a component of $\widehat{P}_{st}^{ko}$, periods of over- or under-valuation are also associated with movements of $\widehat{P}^{ko}_{st}$, providing the necessary correlation. Second, the instrument has to be uncorrelated with the error term ε^{ko}_{st}, 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 κ^{ko}_{st} and μ^{k}_{st} are uncorrelated with the real exchange rate.

We estimate equation (22) using two-stage least squares (2SLS). Our estimate of country k's Quality Index relative to the numeraire is

$$\ln \widehat{\lambda}^{ko}_{st} = -\left(\frac{\widehat{\zeta}^{ko}_{0s} + \widehat{\zeta}^{ko}_{1s} t}{\widehat{\gamma}_s} \right), \tag{23}$$

where t indexes the number of years since 1989 and the remaining right-hand side variables are estimates from equation (22). 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 (22) and the definition of $\ln \widehat{\lambda}^{ko}_{st}$ in equation (23). They are equal to

$$\ln \widehat{\widehat{P}}^{ko}_{st} = \ln \widehat{P}^{ko}_{st} - \ln \widehat{\lambda}^{ko}_{st} = \left(\frac{\widetilde{T}^{k}_{st} - \widehat{\Upsilon}'_{st} - \widehat{\upsilon}^{ko}_{st}}{\widehat{\gamma}_s} \right). \tag{24}$$

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 (ε_{st}^{ko}) are mis-attributed to the Pure Price Index.

V.B..1 Second-Stage Data Requirements

Second-stage data requirements are strong relative to data availability. 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. Further detail is available in our web-based technical appendix.

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.[30] We measure countries' annual net trade in overall manufacturing as well as the industries within manufacturing discussed below. As required by equation (22), we normalize trade balances by total expenditure (E^k). We compute E^k by subtracting total net trade from GDP. Both variables are drawn from the World Bank's World Development Indicators (WDI) database except from Taiwan's GDP, which comes from the Economist Intelligence Unit website. We also need

30. 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 a web-based technical appendix.

to compute the share of manufacturing in total expenditure, $b_s = E_s^k / E^k$. To compute the numerator, we subtract manufacturing net trade from manufacturing value added. The latter variable is drawn from the United Nations' National Accounts Official Country Data. We obtain $b_s = 0.214$ as the average share across countries and years.

We measure trade barriers in terms of transport costs and tariffs. We measure country pairs' bilateral transport costs using 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 $a_{zt}^k = \left(cif_{zt}^k - fob_{zt}^k \right) / fob_{zt}^k$ 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 k on the distance the exports have travelled,

$$\ln a_{zt}^{k,US} = \delta_{st} \ln D^{k,US} + \beta_{st}' X^{k,US} + \in_{zt}^k, \tag{25}$$

where $D^{k,US}$ represents the great circle distance in kilometers between the United States and country k and $X^{k,US}$ represents additional controls, including whether country k 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 $a_{st}^{kk'}$ equal to $\exp\left(\widehat{\delta}_{st} \ln D^{kk'} + \widehat{\beta}_{st}' X^{kk'} \right)$.

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 (22).

We compute bilateral trade costs $\tau_s^{kk'}$ by adding the measures of bilateral transport costs and tariffs explained above. The aggregation of those measures to construct the trade cost term τ_{st}^{k} is more challenging because it requires values for the unobserved consumption price indexes $G_s^{k\prime}$ defined in Section 4. Up to a factor of proportionality (captured by the constant in the regression), the component $\sum_z n_z^{k''} \left(\widetilde{p}_z^{k''}\right)^{1-\sigma_s}$ in the indexes is the share of country k'' in world production of sector s in a world equilibrium with no trade costs. We approximate this share by the share of country k'' 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.[31]

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.

V.B..2 Second-Stage Results

Table III reports second-stage estimates of γ_s from the estimation of equation (22) by OLS and two-stage least squares (2SLS).[32] Robust standard errors

31. The consumption indexes G_s^k also require an estimate of the elasticity of substitution σ_s. We compute τ_{st}^k using $\sigma_s = 6$ and note that alternative values of σ_s ranging from 3 to 10 have almost no impact on our results.

32. 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.

adjusted for clustering at the country level are reported below each coefficient.[33] As indicated in the table, the OLS estimate of γ_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 an order of magnitude lower than the OLS estimate, -0.241 versus -0.028. The final row of the table reports an F-statistic for the first stage of 2SLS of 37.7.

Log Quality Index intercepts and slopes, normalized by annual means as in equation (19), are displayed in Figure III along with their 95 percent confidence bands.[34] Estimated intercepts are equivalent to countries' relative log quality in 1989. As indicated in the figure, China's quality in 1989 is two-thirds ($e^{-0.480}$) that of the mean country in that year, while Germany's is more than twice as high ($e^{0.768}$). Estimated slopes report how much relative quality rises or falls *vis a vis* the mean country each year. Ireland has the highest slope while Hong Kong has the lowest.

Figure III 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. Two noticeable outliers to this pattern are Singapore and Ireland, both of which 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 IV. As indicated in the figure, China's relative quality is flat over time and generally below those of Germany, Japan and Singapore. Relative quality rises for Hun-

33. An estimate and standard error for γ_s that accounts for the fact that the IPIs are estimated are computed using the bootstrap method described in the web-based technical appendix. They are -0.254 and 0.091, respectively. Results are similarly close for our industry-level estimates below.

34. Standard errors are computed using the delta method. Quality intercepts and slopes are reported for each country in tabular form in the web-based technical appendix.

gary, Thailand, Malaysia and Singapore, though estimates for the latter two countries are relatively imprecise.

Our results are robust to a number of sensitivity analyses. First, we obtain similar point estimates after selectively removing each country from the estimation, indicating that results are not overly dependent on the presence of any particular country. Second, we find similar results using more or less stringent secondary screens, though standard errors are generally larger the more lax is the screening. Finally, we assess the potential impact of composition bias by restricting the sample to the 30 largest exporters. This restriction doubles the median number of products country pairs export in common across all years of the sample period, substantially reducing the potential influence of violations to Proposition 1. We find that the rank correlations of relative quality rankings across countries appearing in both samples is above 97 percent in all years.

VI. What Can We Learn from Quality Estimates?

In this section we compare our estimates of countries' manufacturing quality to raw export prices and examine links between quality and long-run growth. We find that changes in raw relative export prices can be a poor approximation of changes in quality, and that consideration of price and quality together provides complementary information about variation in development strategies across countries. Indeed, our results suggest two potential paths towards long-run growth: via quality, as in the case of Malaysia and via price competitiveness, as in the case of China.

One of the most important lessons to take from our estimates of manufacturing quality is that changes in quality inferred from our method can be quite different from changes in countries' relative export prices. Figure V 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 (0.33), 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 many countries such as Argentina, Greece, Portugal, and Switzerland. These divergences between quality and impure prices are due to variation in countries' global net trade balances. Figure VI, for example, shows that Malaysia's rising trade balance combined with its relatively flat impure prices results in rising estimated quality. For China, relatively moderate increases in its manufacturing trade surplus combined with a falling Impure Price Index imply falling pure prices and flat quality.

Quality levels across countries compress over time. The two panels of Table IV display countries' quality rankings and normalized quality levels at four-year intervals from 1989 to 2003. Countries are sorted according to their ranking in 1989, with the final column in each panel noting the change in ranking or level over the entire sample period. Countries whose rank changes by more than ten places are noted with an asterisk. Singapore, Indonesia, Malaysia, and the Philippines are the countries with the largest increases in quality ranking while Australia, Hong Kong, New Zealand, and Poland are those with the largest decreases. This reshuffling of quality rankings is associated with strong quality convergence. The mean quality level of countries below the overall average of all countries rises from -0.48 in 1989 to -0.35 in 2003 while the mean for countries above the overall average increases from 0.42 to just 0.44.[35] A critical question is whether this compression reveals catching up in terms of technological capa-

35. This narrowing is even more dramatic if Ireland is excluded.

bility by the developing world or quality upgrading via the use of higher-quality intermediate inputs without corresponding gains in productivity.

The theory of economic growth has long established a connection between quality and growth. In particular, the "quality ladder" models developed by Grossman and Helpman (1991) and Aghion and Howitt (1992) postulate countries' ability to upgrade quality as a form of productivity gain that also implies increases in GDP per capita. To gain further understanding of the compression in quality levels manifest in our estimates, we compare normalized quality compression to the convergence in exporters' per-capita income over our sample period. We divide our sample countries into two cohorts according to whether per capita GDP is below or above the median in 1989. Interestingly, we find that quality convergence has not been accompanied with convergence in GDP per capita. While the difference between the average quality of high-income versus low-income countries decreased from 0.67 log points in 1989 to 0.38 log points in 2003, this difference remained almost constant for per capita GDP (2.20 log points in 1989 versus 2.14 log points in 2003).

Quality and per capita GDP are strongly correlated in the cross section of countries. The correlation between countries' normalized quality index and their similarly normalized log per capita GDP is positive and statistically significant across all years of our sample, with an average correlation across years of 0.46. However, consistent with the lack of correspondence between quality and per capita income convergence described above, we find that this correlation declines over time, from 0.54 in 1989 to 0.32 in 2003.[36] The weakening association between quality and per capita GDP also appears in the positive but smaller correlation between *changes* in quality and *changes* in per capita income over the sample period (0.30 and significant at the 5 percent level). This correlation is

36. Hummels and Klenow's (2005) estimates of the cross-sectional elasticity of quality with respect to income in 1995 ranges from 0.09 to 0.23.

displayed visually in Figure VII. As indicated in the figure, countries with above average change in quality vary substantially in terms of their per capita GDP growth. For several of these countries, including Ireland, Chile, and Hungary, relatively high quality growth is accompanied by relatively high per capita GDP growth, consistent with a process of quality upgrading based on gains in (quality-adjusted) productivity. Other countries that display substantial quality growth, however, do not exhibit strong growth in income. Malaysia, the Philippines, Indonesia and Thailand, for example, are among the countries with the most impressive increases in manufacturing quality, but this does not appear to have translated in higher incomes. This outcome is consistent with quality growth achieved via use of higher-quality inputs, in which case quality growth might not require enhanced productive capabilities and hence need not raise income.

Figure VII also suggests alternative paths for attaining higher income that are not associated with upgrading quality. Here, China – which combines extraordinary per capita GDP growth with almost no change in quality – serves as perhaps the most interesting and illustrative example of quality growth not being a prerequisite for income growth. There are many potential explanations for this outcome – e.g., the attractiveness of China's large domestic market, its transition from a command economy to a more market-based economy, the effect of its export promotion policies – all of which are worthy of further study.

Overall, the divergence in income and quality growth paths displayed in Figure VII suggests a variety of development strategies that countries might pursue and the usefulness of understanding the economic mechanisms that each strategy involves. Despite its importance, there has been relatively little empirical investigation into the link between product quality in economic development, mostly due to lack of measures of product quality.[37] We hope the method and

37. Hausmann, Hwang and Rodrik (2006), for example, investigate the link between similarity of developing countries' export baskets with those of developed countries and growth.

estimates proposed here help improve this situation.

VII. Conclusion

This paper attempts to fill an important gap in the international trade and development literature by proposing a method for identifying countries' product quality 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., 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' product quality. Further theoretical and empirical efforts on this front will be quite useful.

Estimates of countries' product quality are obviously useful for 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 is scarce. Quality-adjusted price indexes will likely also find use in the labor 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.

Universidad de San Andres and NBER

Yale School of Management and NBER

Appendix I – Theory

Proof of Proposition 1

Since we have already established that $\ln H_s^{kk'} \leq \ln P_s^{kk'} + \ln \phi_{s,c}^{kk'}$, to demonstrate that $\ln H_s^{kk'} \leq \ln P_s^{kk'}$ we only need to show that $\ln \phi_{s,c}^{kk'} \leq 0$, which is equivalent to showing that

$$\sum_z \frac{n_z^k}{\overline{n}_s^k} \Delta \widetilde{p}_z^{kk'} \geq 0. \tag{26}$$

Since $\frac{n_z^k}{\overline{n}_s^k} = \widetilde{n}_z^{k,kk'} + \widetilde{n}_z^{kk'} + \overline{n}_z$, the left-hand-side of (26) can be written as

$$\sum_z \frac{n_z^k}{\overline{n}_s^k} \Delta \widetilde{p}_z^{kk'} = \sum_z \widetilde{n}_z^{k,kk'} \Delta \widetilde{p}_z^{kk'} + \sum_z \widetilde{n}_z^{kk'} \Delta \widetilde{p}_z^{kk'} + \sum_z \overline{n}_z \Delta \widetilde{p}_z^{kk'}. \tag{27}$$

Using the property $\sum\limits_z \widetilde{n}_z^{k,kk'} = 0$, $\sum\limits_z \widetilde{n}_z^{k,kk'} \Delta \widetilde{p}_z^{kk'} = Z_s cov_s \left(\widetilde{n}_z^{k,kk'}, \Delta \widetilde{p}_z^{kk'} \right) \geq 0$ by Assumption 3. Using $\sum\limits_z \widetilde{n}_z^{kk'} = 0$, $\sum\limits_z \widetilde{n}_z^{kk'} \Delta \widetilde{p}_z^{kk'} = Z_s cov_s \left(\widetilde{n}_z^{kk'}, \Delta \widetilde{p}_z^{kk'} \right) = 0$ by Assumption 4. Finally, using the definition of $\widetilde{P}_s^{kk'}$, $\sum\limits_z \overline{n}_z \Delta \widetilde{p}_z^{kk'} = 0$. Combining these results we obtain:

$$\sum_z \frac{n_z^k}{\overline{n}_s^k} \Delta \widetilde{p}_z^{kk'} = Z_s cov_s \left(\widetilde{n}_z^{k,kk'}, \Delta \widetilde{p}_z^{kk'} \right) + Z_s cov_s \left(\widetilde{n}_z^{kk'}, \Delta \widetilde{p}_z^{kk'} \right) \geq 0,$$

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

$$\ln H_s^{kk'} \leq \ln P_s^{kk'} \leq \ln L_s^{kk'}.$$

Proof of Proposition 2

Substituting $n_z^k = \overline{n}_s^k(\overline{n}_z + \widetilde{n}_z^k)$ into equation (14), we can rewrite the right-hand side of that equation as

$$-1 + \frac{\overline{n}_s^k}{E^k}\left(\widetilde{P}_s^k\right)^{1-\sigma_s}\left(\exp\tau_s^k\right)\left[\sum_z \overline{n}_z\left(\frac{\widetilde{p}_z^k}{\widetilde{P}_s^k}\right)^{1-\sigma_s} + \sum_z \widetilde{n}_z^k\left(\frac{\widetilde{p}_z^k}{\widetilde{P}_s^k}\right)^{1-\sigma_s}\right].$$

Using the definition of $\widetilde{P}_s^k$ and the fact that $\sum_z \widetilde{n}_z^k(\frac{\widetilde{p}_z^k}{\widetilde{P}_s^k})^{1-\sigma_s} = Z_s cov_s\left(\widetilde{n}_z^k, (\frac{\widetilde{p}_z^k}{\widetilde{P}_s^k})^{1-\sigma_s}\right)$, the above expression can be rewritten as

$$-1 + \frac{\overline{n}_s^k}{Y^k}\frac{Y^k}{E^k}\left(\widetilde{P}_s^k\right)^{1-\sigma_s}\left(\exp\tau_s^k\right)\left[1 + Z_s cov_s\left(\widetilde{n}_z^k, \left(\frac{\widetilde{p}_z^k}{\widetilde{P}_s^k}\right)^{1-\sigma_s}\right)\right].$$

Using Assumption 5, equation (15), and the fact that $\frac{Y^k}{E^k} = 1 + \frac{T^k}{E^k}$, we can substitute the latter expression for the right-hand side of (14). Rearranging terms and taking natural logarithms, we obtain

(28)
$$\ln\left(1 + \frac{T_s^k}{b_s E^k}\right) = \ln\left[\left(1 + \frac{T^k}{E^k}\right)\left(\widetilde{P}_s^k\right)^{1-\sigma_s-\eta_s}\left(\exp\tau_s^k\right)\left[1 + Z_s\left(\varphi_s + \mu_s^k\right)\right]\right].$$

Using $\ln(1+x) \simeq x$, and abstracting from the approximation error, we can finally express equation (28) as

$$\frac{T_s^k - b_s T^k}{E^k} = \Upsilon_s + \gamma_s \ln\widetilde{P}_s^k + b_s\tau_s^k + \iota_s^k$$

where

$$\Upsilon_s = b_s Z_s \varphi_s, \quad \gamma_s = b_s\left(1 - \sigma_s - \eta_s\right) < 0, \quad \iota_s^k = b_s Z_s \mu_s^k$$

Imperfect Overlap in Active Products

Proposition 1 assumes that all countries export the same products. Here, we outline the more general case in which the set of exported products varies across countries. Let I_s be the set of all product categories in sector s. Define a country as "active" in product z if it reports positive exports in that category. The set I_s can be decomposed into two parts: $I_{s(kk')}$, which includes categories in which countries k and k' are both active (with $Z_{s(kk')}$ denoting the number of such categories), and its complement, $I^{-}_{s(kk')}$. Accordingly, define $\mathbf{p}^{k''}_{s(kk')}$ and $\mathbf{q}^{k''}_{s(kk')}$ as the vector of prices and quantities, respectively, of imports from country $k'' \in \{k, k'\}$ in product categories included in $I_{s(kk')}$. The vector $\mathbf{q}^{-k}_{s(kk')}$ is the complement of $\mathbf{q}^{k}_{s(kk')}$ with respect to $\mathbf{q}$.

In this setting, the constrained expenditure function should be defined over product categories in $I_{s(kk')}$. As a result, $m_{s,k(kk')}(\mathbf{p}^{k''}_{s(kk')}, \mathbf{q}^{-k}_{s(kk')}, \mathbf{n}, \boldsymbol{\lambda}, \boldsymbol{\xi}, u)$ solves the problem

(29)

$$\min_{\widehat{q}^{k}_{s(kk')}} \; \mathbf{p}^{k''}_{s(kk')}\widehat{\mathbf{q}}^{k}_{s(kk')} \qquad s.t. \quad U(\widehat{\mathbf{q}}^{k}_{s(kk')}, \mathbf{q}^{-k}_{s(kk')}, \mathbf{n}, \boldsymbol{\lambda}, \boldsymbol{\xi}) = u, \qquad k'' = 1, ..., K.$$

The solution to problem (29) yields an expression analogous to (5) but with summations ranging over elements of $I_{s(kk')}$. Similarly, inequalities (7) and (9) continue to hold as long as $\phi^{kk'}_{s,k}$ and $\phi^{kk'}_{s,k'}$ are redefined over products in $I_{s(kk')}$.

For the Impure Price Index to be bounded from below by the Paasche Index, it must be the case that $\sum_{z \in I_{s(kk')}} \frac{n^k_z}{\overline{n}^k_s} \Delta\widetilde{p}^{kk'}_z \geq 0$. Given that $\frac{n^k_z}{\overline{n}^k_s} = \widetilde{n}^{k,kk'}_z + \overline{n}^{kk'}_z$, and that $I_{s(kk')}$ is the complement of $I^{-}_{s(kk')}$ with respect to I_s, we can write

(30)

$$\sum_{z \in I_{s(kk')}} \frac{n^k_z}{\overline{n}^k_s} \Delta\widetilde{p}^{kk'}_z = \sum_{z \in I_{s(kk')}} \widetilde{n}^{k,kk'}_z \Delta\widetilde{p}^{kk'}_z + \sum_{z \in I_s} \overline{n}^{kk'}_z \Delta\widetilde{p}^{kk'}_z - \sum_{z \in I^{-}_{s(kk')}} \overline{n}^{kk'}_z \Delta\widetilde{p}^{kk'}_z .$$

The first term in (30) can be expressed as

$$
\begin{aligned}
(31) \sum_{z\in I_{s(kk')}} \widetilde{n}_z^{k,kk'}\Delta\widetilde{p}_z^{kk'} &= Z_{s(kk')}cov_{I_{s(kk')}}\left(\widetilde{n}_z^{k,kk'},\Delta\widetilde{p}_z^{kk'}\right)+\overline{\Delta\widetilde{p}_s^{kk'}}\sum_{z\in I_{s(kk')}}\widetilde{n}_z^{k,kk'} \\
&\geq -\overline{\Delta\widetilde{p}_s^{kk'}}\sum_{z\in I^-_{s(kk')}}\widetilde{n}_z^{k,kk'}
\end{aligned}
$$

where the second line uses Assumption 3 – now only including products in $I_{s(kk')}$ in the covariance – and the property $\sum_{z\in I_{s(kk')}}\widetilde{n}_z^{k,kk'} = -\sum_{z\in I^-_{s(kk')}}\widetilde{n}_z^{k,kk'}$, and where $\overline{\Delta\widetilde{p}_s^{kk'}} \equiv \frac{1}{Z_{s(kk')}}\sum_{z\in I_{s(kk')}}\Delta\widetilde{p}_z^{kk'}$. The second term in (30) has already been shown to equal zero in the proof of Proposition 1 (note that $\overline{n}_z^{kk'}=\widetilde{n}_z^{kk'}+\overline{n}_z$).

Combining these results we obtain

$$
\sum_{z\in I_{s(kk')}}\frac{n_z^k}{\overline{n}_s^k}\Delta\widetilde{p}_z^{kk'} \geq -\sum_{z\in I^-_{s(kk')}}\left(\widetilde{n}_z^{k,kk'}\overline{\Delta\widetilde{p}_s^{kk'}}+\overline{n}_z^{kk'}\Delta\widetilde{p}_z^{kk'}\right),
$$

which is greater than zero (i.e., Impure Price Index is bounded from below by the Paasche Index) only if

$$
(32) \qquad \sum_{z\in I^-_{s(kk')}}\left(\widetilde{n}_z^{c,kk'}\overline{\Delta\widetilde{p}_s^{kk'}}+\overline{n}_z^{kk'}\Delta\widetilde{p}_z^{kk'}\right)\leq 0.
$$

Analogously, to guarantee that the Impure Price Index is bounded above by the Laspeyres Index, we need

(33)
$$
\sum_{z\in I^-_{s(kk')}}\left(\widetilde{n}_z^{k',kk'}\overline{\Delta\widetilde{p}_s^{k'k}}+\overline{n}_z^{kk'}\Delta\widetilde{p}_z^{k'k}\right)=\sum_{z\in I^-_{s(kk')}}\left(\widetilde{n}_z^{k,kk'}\overline{\Delta\widetilde{p}_s^{kk'}}-\overline{n}_z^{kk'}\Delta\widetilde{p}_z^{kk'}\right)\leq 0.
$$

A sufficient condition for (32) and (33) to hold simultaneously is that the two countries are active in the same set of goods – the assumption made in Section 3 to simplify the exposition. In that case the set $I^-_{s(kk')}$ is empty and the left hand side of both inequalities are zero since they sum over elements of an empty set. Unfortunately, since countries often are not "active" in the same set of products, we need to analyze how imperfect overlap in active products might violate the result of Proposition 1.

The summations in conditions (32) and (33) include two terms. The first term ($\widetilde{n}_z^{c,kk'}\overline{\Delta\widetilde{p}_s^{kk'}}$) is common to both conditions while the second term enters each condition with the opposite sign. As a result, the larger the absolute value of the second term, the more likely one of the two conditions is violated.

In considering the absolute magnitude of the second term, note that it is a weighted sum of $\Delta\widetilde{p}_z^{kk'}$ (defined in equation (11)) across "mismatched" products, i.e. products in which one country is active but the other one is not. Also note that $\Delta\widetilde{p}_z^{kk'} > 0$ when k is active and k' is not, and *vice versa*. Thus, other things equal, the absolute magnitude of this term increases with the number of "mismatched" categories in the country pair and the asymmetry with which these mismatched products are distributed across countries in the pair. Violations are less likely, for example, when countries k and k' are each active in ten products not produced by the other than when c is active in twenty products not produced by k' but every product produced by k' is also produced by k. In the former, the elements of the sum will tend to cancel out as they have opposite signs. In the latter, condition (32) is not satisfied due to composition: while the Impure Price Index $P_s^{kk'}$ is defined over I_s, the Paasche index we observe, $H_{s(kk')}^{kk'}$, is computed over the subset $I_{s(kk')}$. Therefore, even though $H_s^{kk'} \leq P_s^{kk'}$ is true, $H_{s(kk')}^{kk'} > P_s^{kk'}$.[38] Put another way, in this example, the subset $I_{s(kk')}$ is not a representative sample of the broader set I_s over which $P_s^{kk'}$ is defined. As a result, the Paasche index fails to include (mismatched)

38. Since $p_z^{k'} = \infty$ for any non-active z, including those products in the Paasche index would result in $\ln H_s^{kk'} = -\infty$, as the corresponding elements $p_z^{k'} q_z^k$ in denominator of the index will be infinite. Thus, $H_s^{kk'} \leq P_s^{kk'}$ will be (trivially) true but not informative to estimate $P_s^{kk'}$.

products that have a particularly low price in k relative to the (infinite) price in k'.[39]

The first term in conditions (32) and (33) is less problematic. We can write

$$\overline{\Delta\widetilde{p}_s^{kk'}} = \frac{1}{Z_{s(kk')}} \left(\sum_{z \in I_s} \Delta\widetilde{p}_z^{kk'} - \sum_{z \in I^-_{s(kk')}} \Delta\widetilde{p}_z^{kk'} \right). \tag{34}$$

It is not possible to sign $\sum_{z \in I_s} \Delta\widetilde{p}_z^{kk'}$. However, as a benchmark we can use the fact that this sum is similar – except for the absence of weights – to $\sum_{z \in I_s} \overline{n}_z \Delta\widetilde{p}_z^{kk'}$, which we have shown equals 0. Thus, its departure from zero will depend on the extent to which substituting a constant weight of 1 for $\overline{n}_z$ affects the sum.

Abstracting from this sum, the first term in (32) and (33) can be written as

$$-\frac{1}{Z_{s(kk')}} \left(\sum_{z \in I^-_{s(kk')}} \widetilde{n}_z^{k,kk'} \right) \left(\sum_{z \in I^-_{s(kk')}} \Delta\widetilde{p}_z^{kk'} \right). \tag{35}$$

If mismatched products are asymmetrically active in country k, then both terms in parentheses in equation (35) are positive and the expression overall is negative, as needed to satisfy both conditions simultaneously. If mismatched products are asymmetrically active in country k', then both parentheses are negative but expression (35) is still negative. In sum, even though it is not possible to demonstrate the sign of the first term in (32) and (33), this analysis suggests that it is likely to be negative and thus help satisfy both conditions.

In our empirical analysis, we attempt to mitigate violations of conditions (32) and (33) by excluding country pairs with very few (i.e., less than 25) export products in common. As a robustness check, we have also re-estimated quality for the 30 largest countries in the sample, whose export-product overlap

39. The Laspeyres index, by constrast, is unaffected since, as $q_z^{k'} = 0$ for non-active products in k', including those products only adds zeros to its numerator and denominator.

is substantially higher than for the larger sample. As noted in the main text, quality rankings across those 30 countries are very highly correlated with the rankings reported in our baseline estimation.

Appendix II – Estimating Quality Within Manufacturing

In this appendix we compute separate Quality Indexes for the four one-digit SITC industries that constitute manufacturing. To explore the potential influence of countries' use of intermediate inputs outside of the sectors at which quality is being estimated, we also estimate quality across the two, two-digit SITC sectors that represent apparel and textiles.

One-digit SITC Sectors

Estimation of 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 A.1 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 A.1 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 A.2 reports the second-stage OLS (top panel) and 2SLS (bottom panel) estimates of γ_s by industry.[40] The 2SLS estimates of γ_s have the expected negative sign in all four industries, but the estimate for Chemicals is 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 A.1).

Normalized log Quality Index intercepts and slopes along with their standard errors are displayed along with their 95 percent confidence bands in Figure A.2[41]

40. We compute b_s for each one-digit sector using the procedure for overall manufacturing described above. The values are 0.035, 0.043, 0.072, and 0.034 for SITC 5, 6, 7, and 8, respectively. For more detail, see our technical appendix.

41. Standard errors are computed using the delta method. Quality intercepts and slopes are reported for each country and industry in tabular form in the web-based technical appendix.

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 relatively high in Manufactured Materials and Miscellaneous Manufactures, which include Textiles (SITC 65) and Apparel (SITC 84), respectively. Given the weak results for the Chemicals sector, we exclude if from further analysis.

Disaggregated quality estimates reveal substantial variation in quality intercepts across industries within countries. China, for example, is at the low end in both Manufactured Materials and Machinery but in the middle of the pack in Miscellaneous Manufacturing. Hong Kong and Taiwan, on the other hand, have relatively high intercepts for Miscellaneous Manufactures but are in the bottom tercile of Machinery. Overall, we find a positive and statistically significant correlation of quality intercepts across countries for Manufactured Materials and Machinery but little correlation between Miscellaneous Manufactures and the other two sectors.

Quality Index slopes display similar variation: across countries the non-Chemical industry slopes have the same sign in only 14 of the 43 countries in the sample. These differences are highlighted in Figure A.3 for the subset of nine countries examined in Figure IV. For Singapore, relative quality increases strongly in all three sectors. For Malaysia, quality increases strongly in Machinery but is relatively flat in Manufactured Materials and Miscellaneous Manufactures.

Quality convergence within manufacturing also varies across industries. Figure A.4 reports the evolution of mean quality for countries with initially high and low per capita income. Trends are displayed for overall manufacturing as well as for Manufactured Materials, Machinery and Miscellaneous Manufactures. 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.

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 effect 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 either value added trade data or input-output tables to map imported intermediates to final goods. Unfortunately, data for pursing these solutions is generally unavailable across countries and time. Here, we take advantage of the well-known linkage between Apparel (SITC 84) and Textiles (SITC 65) to examine separately the estimated quality of apparel versus a hybrid sector, Apparel & Textiles (SITC 65&84), that combines the goods from both two-digit SITC industries.

Table A.3 reports 2SLS estimates of γ_s for SITC 65, SITC 84 and the hybrid Apparel & Textiles. As indicated in the table, estimates of γ_s are negative and significant for all three groups of goods.[42] Figure A.5 compares normal-

42. Quality intercepts and slopes for all three sets of goods are reported for each country in tabular form in the web-based technical appendix.

ized Apparel versus Apparel & Textiles quality across countries in 2003. While estimated quality for the two sets of goods is highly correlated, outliers are apparent. Pakistan, for example, generally has a higher trade surplus in textiles than apparel. As a result, normalized quality for Apparel & Textiles is substantially higher than it is for Apparel alone.

Such outliers suggest that controlling for intermediate inputs may have an important influence on estimated quality. As a result, it seems prudent to include as much information about input-output linkages as possible when estimating quality across disaggregate product categories. Over time, as collection and dissemination of more detailed data on countries' international trade becomes available, this task should become easier.

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Table I
Sample Attributes for All Manufacturing in 2003, by Screening

	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*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 II
Sample and First-Stage Estimation Attributes, All Manufacturing

	Attributes of Estimation Sample					Attributes of First-Stage Estimation		
	Countries	Country-Pair Observations	Median Common Products Across Country Pairs	Product-Country-Pair Observations	Median Log Paasche-Laspeyres Interval	Objective Function	Maximization Standard Error	First-Stage Estimates Within Bounds
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 III
Second-Stage Regression Results for All Manufacturing

	OLS	2SLS
Impure Price Index	-0.028	-0.241 ***
	0.023	0.084
Observations	640	640
R^2	0.93	0.90
First-Stage F Statistic	.	37.7

Notes: Table displays OLS and 2SLS regression results for estimation of equation (22) 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 IV
Quality Rankings, All Manufacturing

Country	Rank 1989	1993	1998	2003	Change	Normalized Quality 1989	1993	1998	2003	Change
Switzerland (CHE)	1	2	2	4	-3	0.93	0.84	0.73	0.62	-0.31
Sweden (SWE)	2	3	5	6	-4	0.83	0.75	0.65	0.55	-0.28
Germany (DEU)	3	5	7	9	-6	0.77	0.66	0.54	0.41	-0.36
Finland (FIN)	4	4	3	3	1	0.67	0.67	0.67	0.67	0.00
Italy (ITA)	5	6	8	8	-3	0.66	0.59	0.51	0.42	-0.24
France (FRA)	6	8	9	10	-4	0.63	0.54	0.44	0.34	-0.29
Japan (JPN)	7	9	10	12	-5	0.57	0.47	0.33	0.20	-0.38
Belgium (BEL)	8	7	6	5	3	0.54	0.55	0.55	0.55	0.01
United Kingdom (GBR)	9	10	12	17	-8	0.47	0.38	0.26	0.15	-0.33
Denmark (DNK)	10	11	11	14	-4	0.45	0.37	0.27	0.17	-0.28
Ireland (IRL)	11	1	1	1	10	0.45	0.87	1.41	1.94	1.50
Austria (AUT)	12	12	13	16	-4	0.42	0.34	0.25	0.15	-0.27
Israel (ISR)	13	13	16	19	-6	0.38	0.27	0.14	0.02	-0.36
*Australia (AUS)	14	19	23	26	-12	0.27	0.11	-0.10	-0.31	-0.58
Taiwan (TWN)	15	17	18	22	-7	0.24	0.15	0.03	-0.09	-0.33
Spain (ESP)	16	18	19	24	-8	0.23	0.14	0.02	-0.10	-0.33
*Hong Kong (HKG)	17	21	25	31	-14	0.23	0.06	-0.15	-0.36	-0.59
Canada (CAN)	18	22	24	27	-9	0.21	0.06	-0.13	-0.31	-0.52
Norway (NOR)	19	20	20	21	-2	0.18	0.10	0.01	-0.08	-0.26
Netherlands (NLD)	20	16	15	13	7	0.17	0.17	0.17	0.18	0.01
Korea, Republic of (KOR)	21	15	14	11	10	0.17	0.19	0.21	0.23	0.06
*New Zealand (NZL)	22	25	30	38	-16	0.08	-0.08	-0.28	-0.48	-0.57
Portugal (PRT)	23	23	22	20	3	0.01	0.00	-0.01	-0.02	-0.04
Argentina (ARG)	24	26	26	25	-1	-0.11	-0.14	-0.18	-0.21	-0.10
Hungary (HUN)	25	24	17	15	10	-0.16	-0.07	0.04	0.16	0.31
Brazil (BRA)	26	27	29	28	-2	-0.19	-0.23	-0.28	-0.33	-0.14
*Singapore (SGP)	27	14	4	2	25	-0.19	0.19	0.66	1.13	1.31
Mexico (MEX)	28	28	31	33	-5	-0.33	-0.35	-0.38	-0.41	-0.08
Turkey (TUR)	29	29	33	34	-5	-0.33	-0.37	-0.41	-0.45	-0.11
Greece (GRC)	30	30	34	35	-5	-0.41	-0.42	-0.43	-0.45	-0.03
Romania (ROM)	31	32	38	40	-9	-0.42	-0.45	-0.48	-0.51	-0.09
*Poland (POL)	32	35	41	43	-11	-0.42	-0.47	-0.53	-0.59	-0.17
Colombia (COL)	33	36	40	41	-8	-0.45	-0.47	-0.50	-0.52	-0.07
South Africa (ZAF)	34	31	32	29	5	-0.46	-0.42	-0.38	-0.35	0.11
China (CHN)	35	37	37	37	-2	-0.48	-0.48	-0.48	-0.48	0.00
India (IND)	36	38	36	36	0	-0.52	-0.50	-0.48	-0.45	0.07
*Indonesia (IDN)	37	33	28	23	14	-0.59	-0.45	-0.27	-0.09	0.50
Morocco (MAR)	38	40	35	30	8	-0.60	-0.53	-0.44	-0.35	0.25
Thailand (THA)	39	41	39	32	7	-0.68	-0.59	-0.48	-0.37	0.31
Pakistan (PAK)	40	42	42	39	1	-0.69	-0.63	-0.56	-0.49	0.20
*Philippines (PHL)	41	39	27	18	23	-0.74	-0.52	-0.24	0.04	0.78
*Malaysia (MYS)	42	34	21	7	35	-0.83	-0.46	0.01	0.47	1.31
Chile (CHL)	43	43	43	42	1	-0.95	-0.84	-0.71	-0.58	0.37

Notes: Table records countries' quality ranking and normalized quality indexes by year. Countries are sorted according to their 1989 ranking. Final column of each panel notes the change between 1989 and 2003. * denotes countries whose rank changes by more than ten places between 1989 and 2003.

Table A.1

First-Stage Optimization Results, by Manufacturing Industry

	First-Stage Estimates Within Bounds	Objective Function	Maximization Standard Error	Median Log Paasche-Laspeyres Interval	Product-Country-Pair Obs	Country Pair Obs	Median Common Products Across Country Pairs	First-Stage Estimates Within Bounds	Objective Function	Maximization Standard Error	Median Log Paasche-Laspeyres Interval	Product-Country Pair Obs	Country Pair Obs	Median Common Products Across Country Pairs
	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 A.2

Second-Stage Regression Results, by Manufacturing Industry

	OLS			
	5 - Chemicals	6 - Manuf Mat	7 - Machinery	8 - Misc Manuf
Impure Price Index	0.007	-0.002	-0.015 ***	0.003
	0.008	0.004	0.007	0.006
Observations	400	608	533	610
R^2	0.97	0.97	0.92	0.96

	2SLS			
	5 - Chemicals	6 - Manuf Mat	7 - Machinery	8 - Misc Manuf
Impure Price Index	0.089	-0.171 ***	-0.090 ***	-0.055 *
	0.130	0.085	0.041	0.037
Observations	400	608	533	610
R^2	0.91	0.77	0.89	0.94
First-Stage F Statistic	0.01	4.21	34.3	13.6

Notes : Table reports OLS and 2SLS regression results for equation (22). 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 A.3
Second-Stage Regression Results for Apparel and Textiles

	2SLS		
	SITC 65	SITC 84	SITC 65&84
Impure Price Index	-0.022	-0.054 *	-0.061 *
	0.017	0.031	0.034
Observations	434	528	579
R^2	0.97	0.91	0.95
First-Stage F Statistic	4.6	11.8	16.0

Notes: Table compares 2SLS regression results for estimation of equation (22) 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.

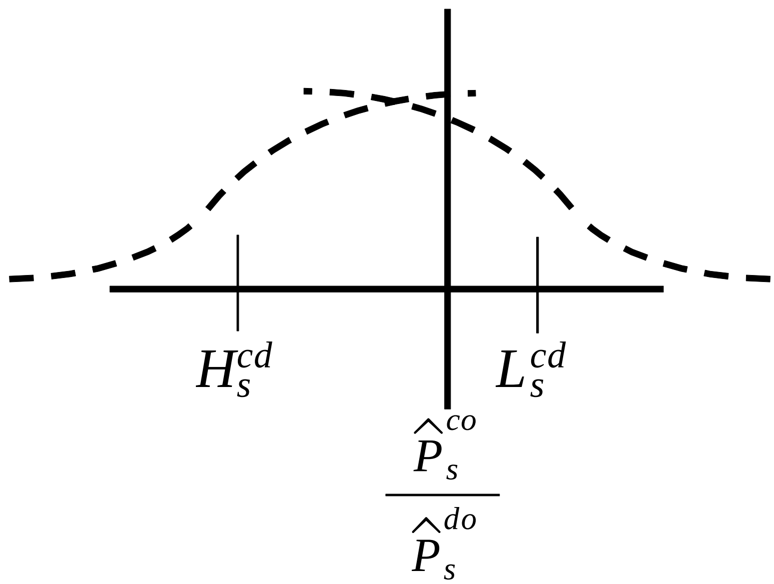

Figure I
Maximizing the Likelihood that the Observed Paasche and
Laspeyres Bounds Contain the Estimated Impure Price Index

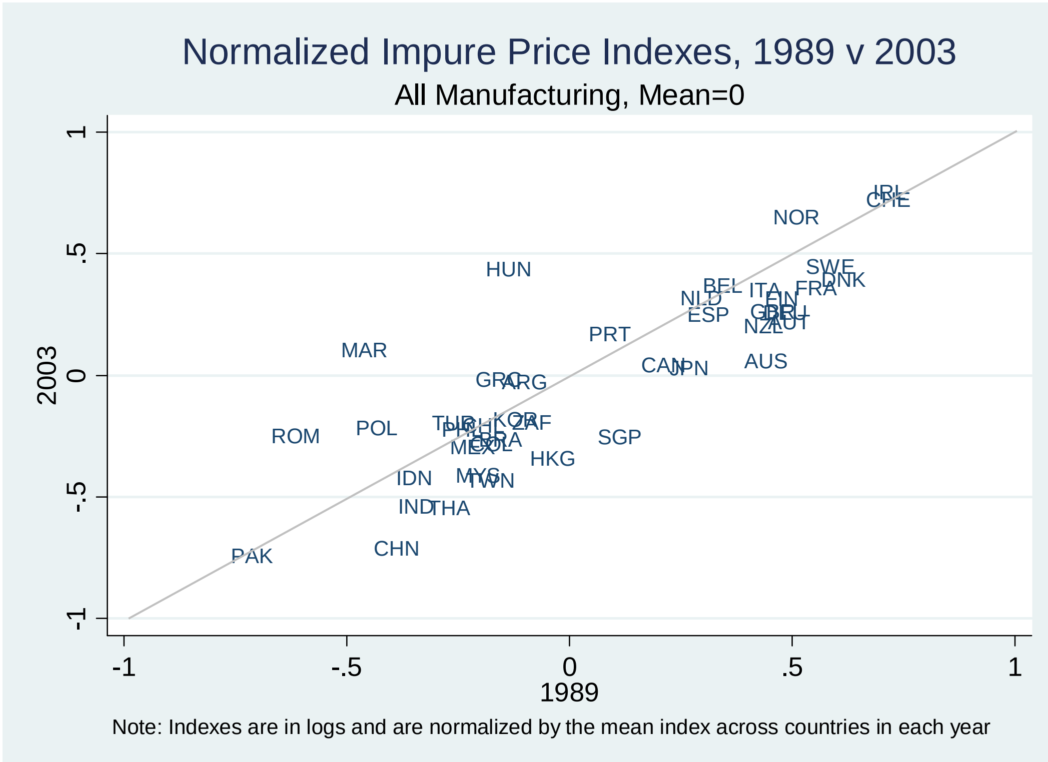

Figure II
First-Stage Estimated Impure Price Indexes, 2003 Versus 1989

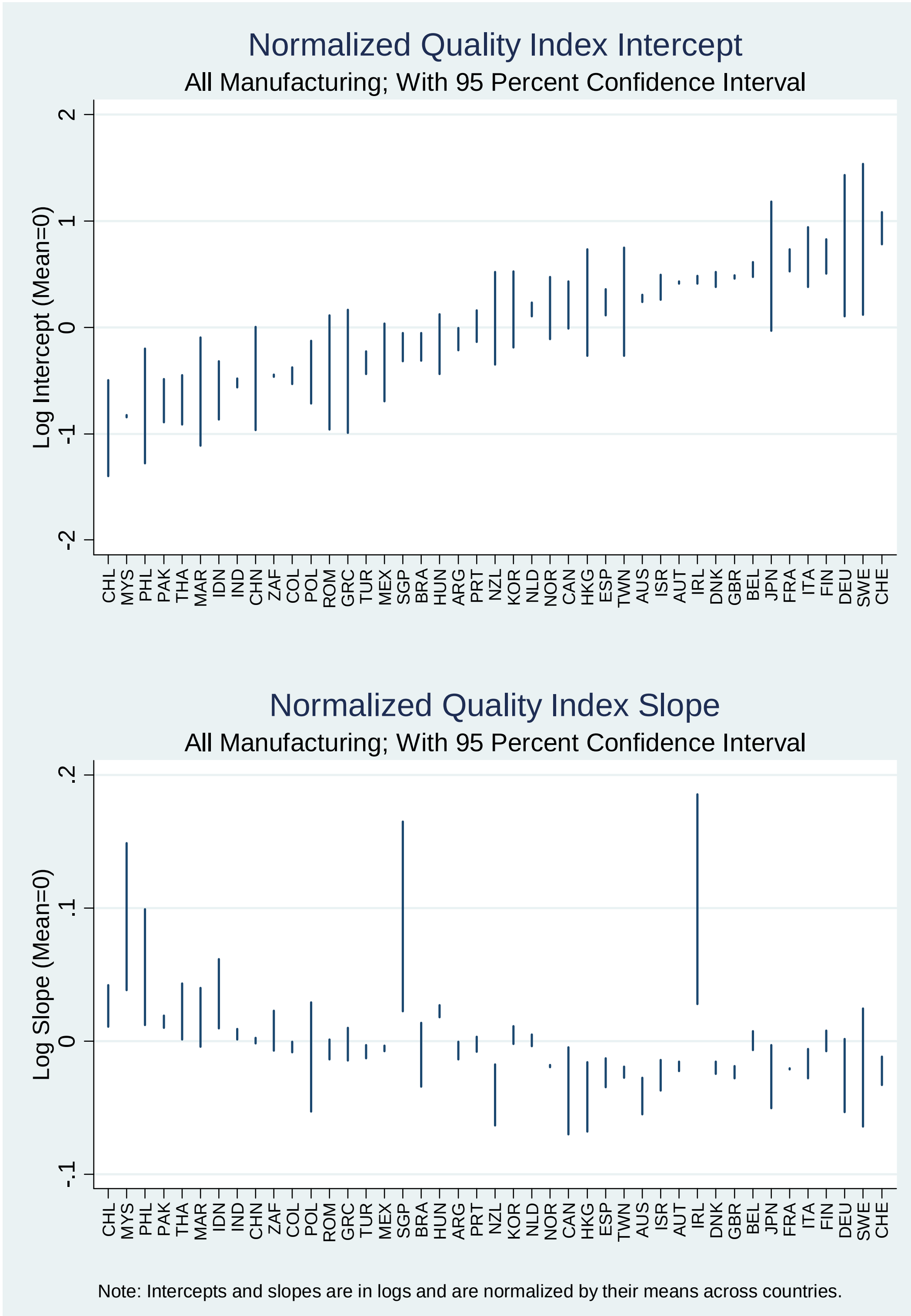

Figure III

Normalized Log Quality Index Intercept and Slope, by Country

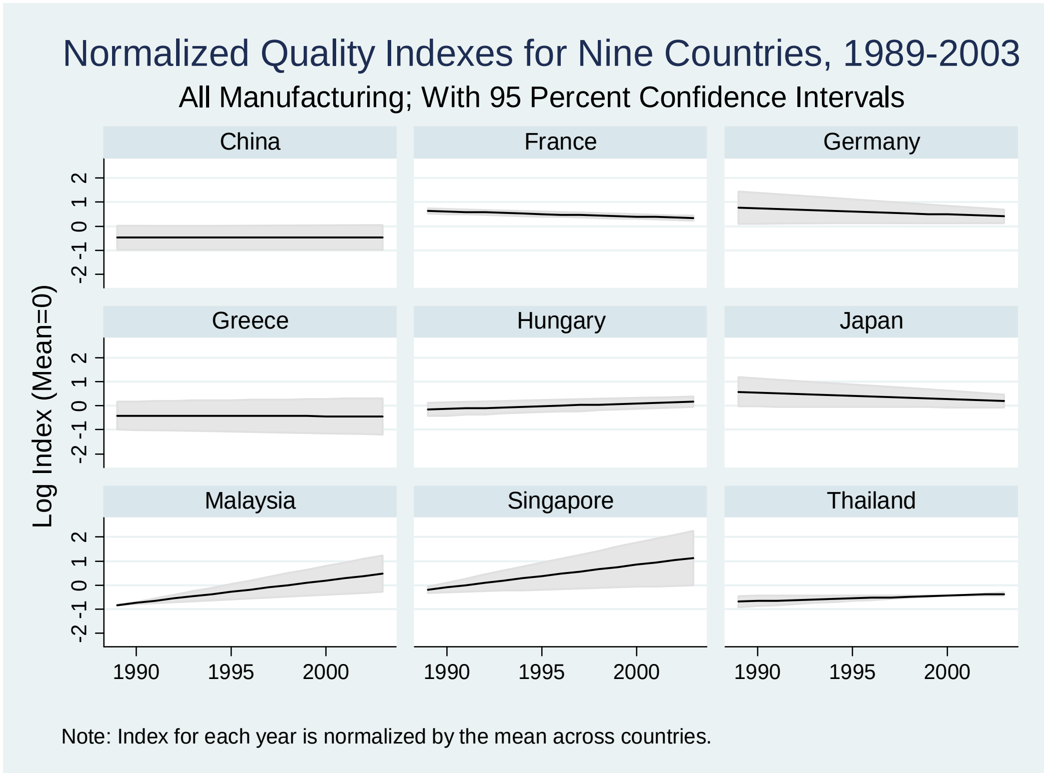

Figure IV
Normalized Log Quality Index for Select Countries, 1989 to 2003

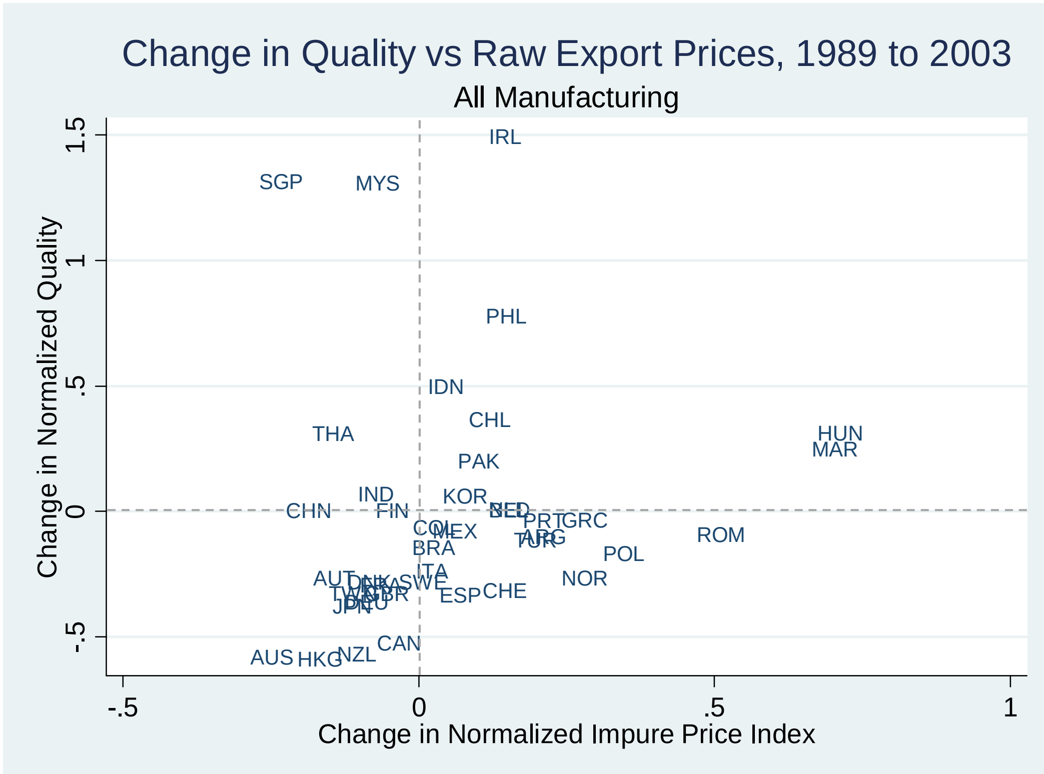

Figure V

Normalized Quality Versus Change in Normalized Impure Price Index

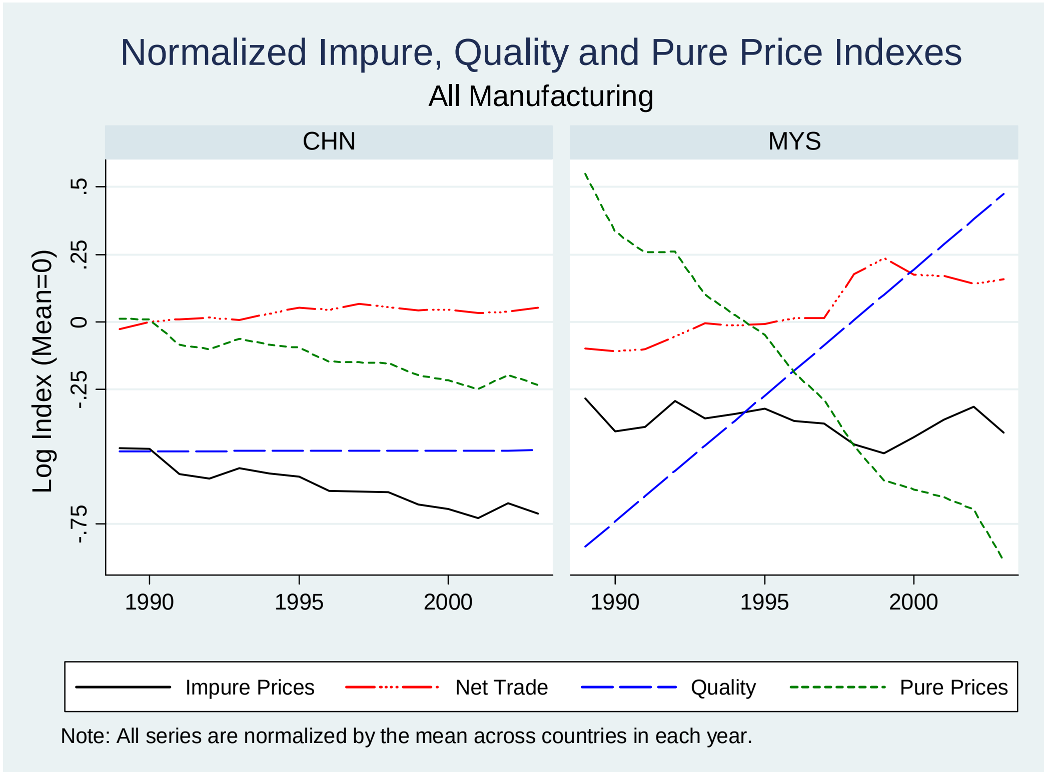

Figure VI
Decomposition of China's and Malaysia's Impure Price Index

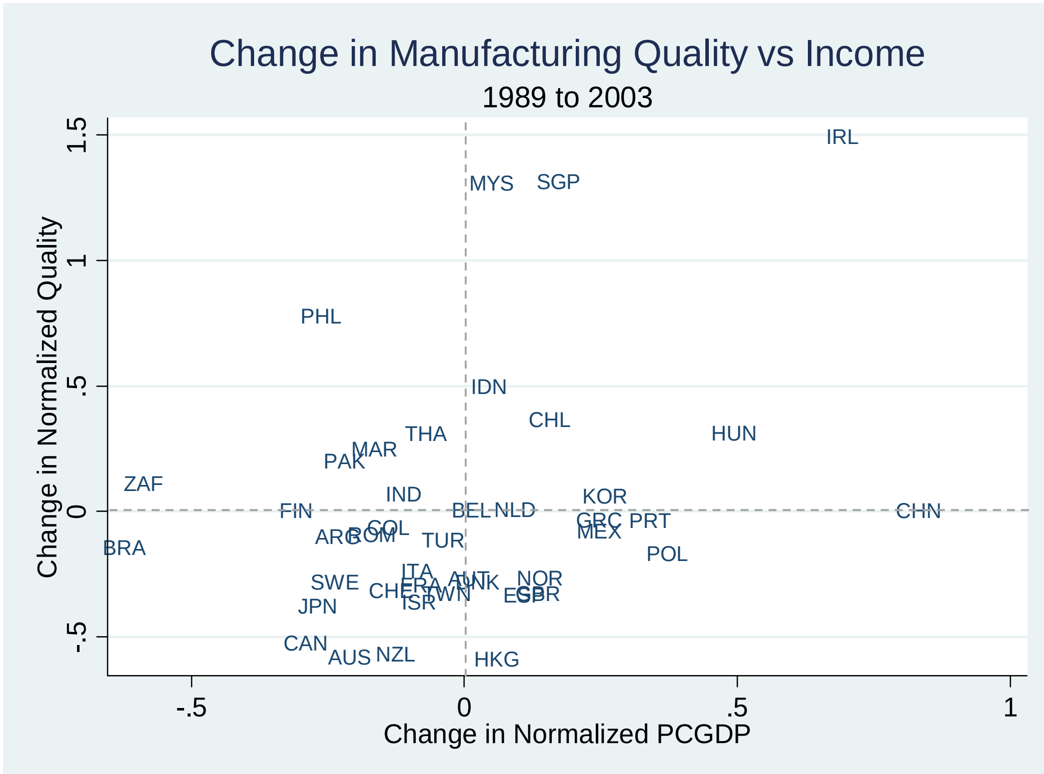

Figure VII

Change in Normalized Quality Versus Change in Normalized Income

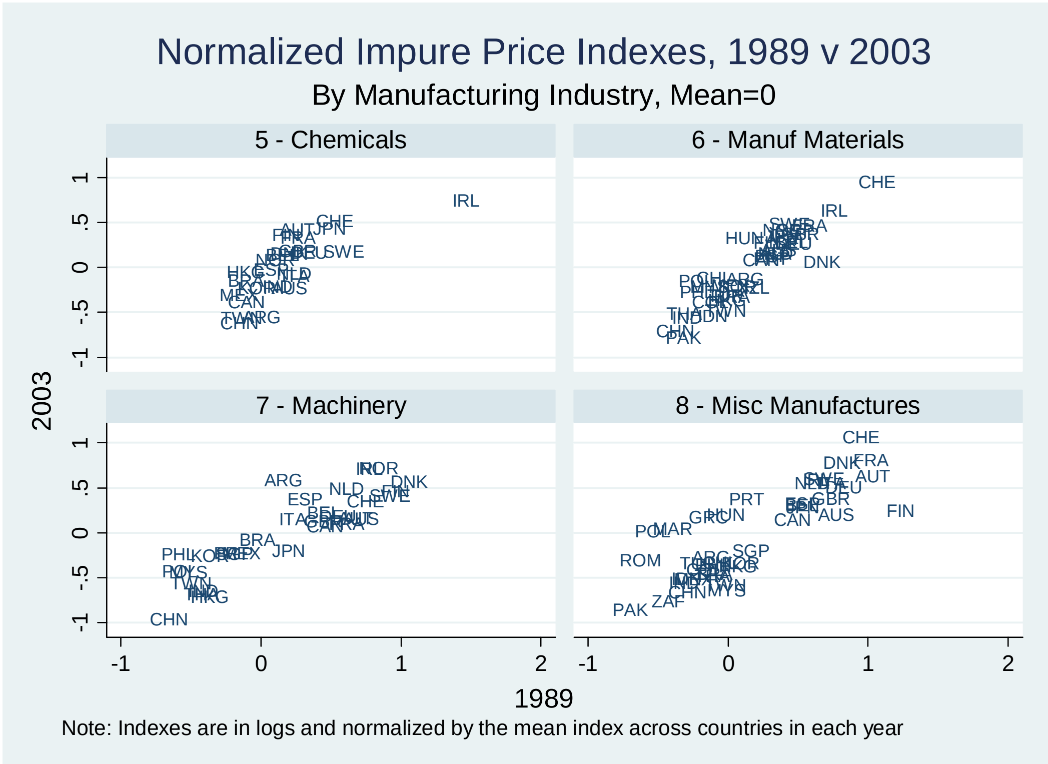

Figure A.1
Normalized Impure Price Indexes

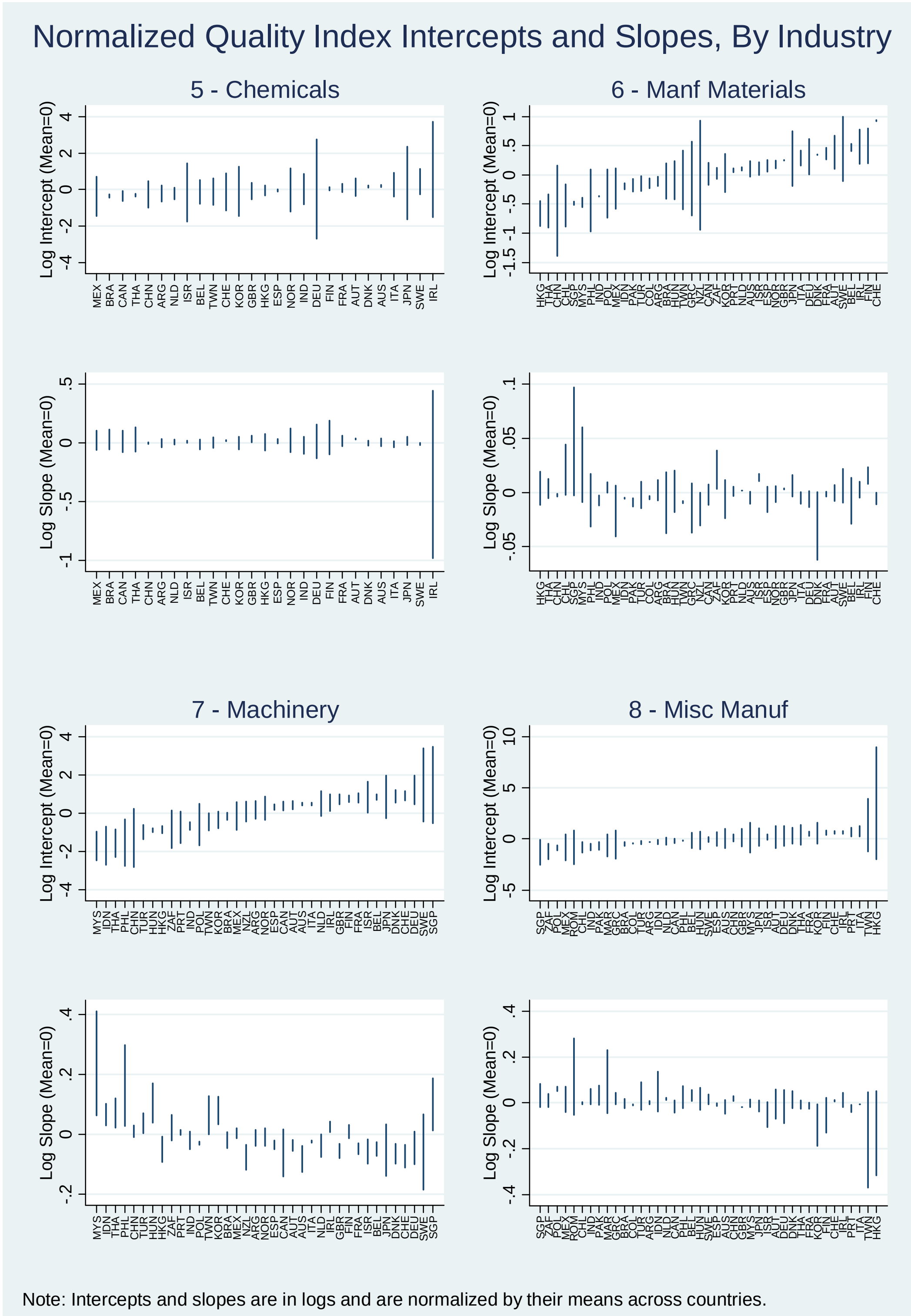

Figure A.2
Normalized Log Quality Index Intercepts and Slopes, by Manufacturing Industry

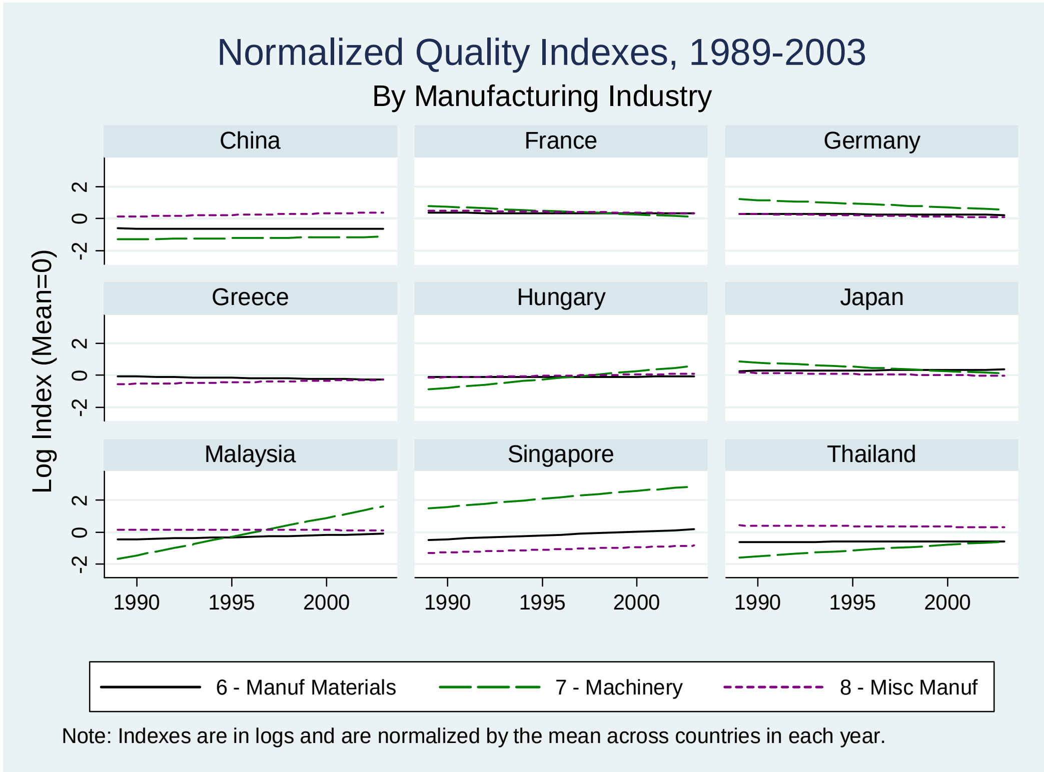

Figure A.3

Normalized Log Quality Indexes For Select Countries, by Manufacturing Industry

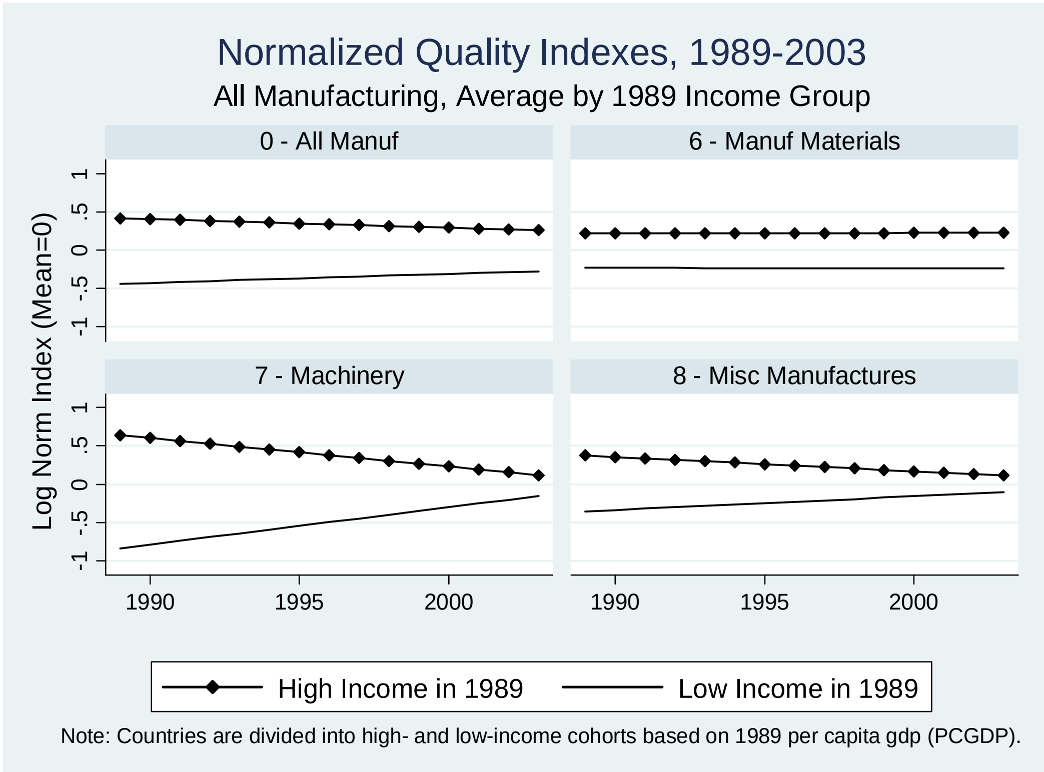

Figure A.4

Evolution of Mean Normalized Quality for Countries with High and Low Income in 1989

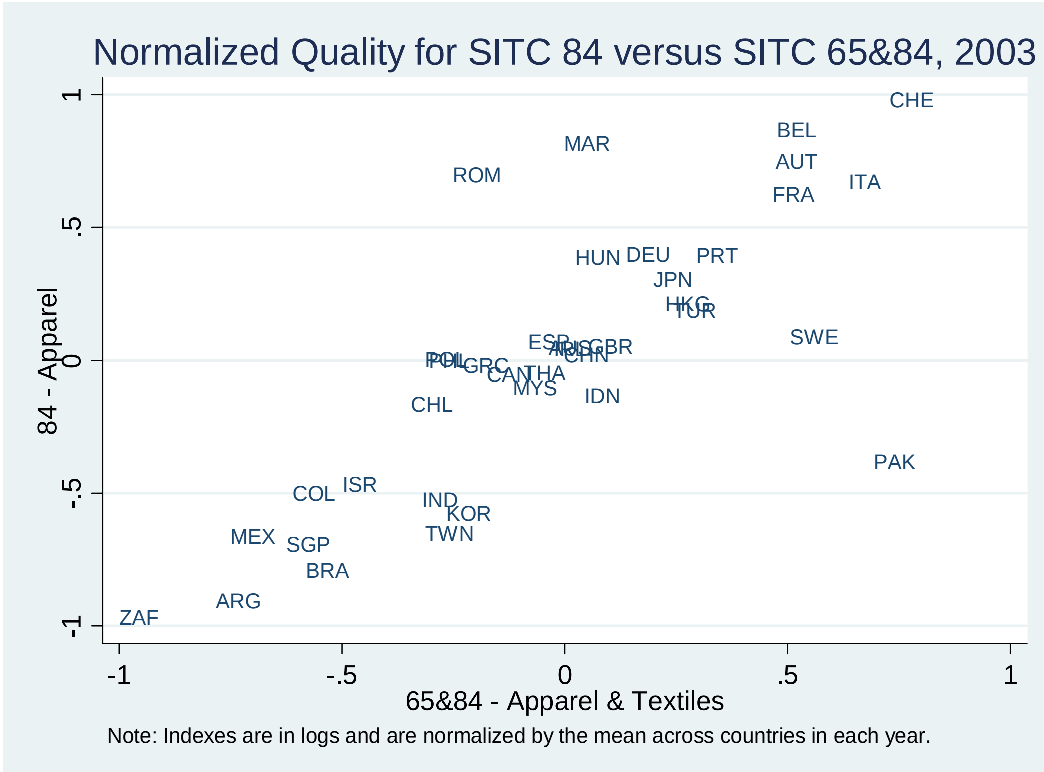

Figure A.5

Comparison of 2003 Quality Indexes for Apparel Versus Apparel & Textiles