Technical Appendix for "Estimating Cross-Country Differences in Product Quality"*

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Abstract

This web-based technical appendix provides additional information on the data and estimation techniques used in the above-referenced paper. It also reports the full set of quality index intercepts and slopes discussed in the paper in tabular form, as well as their standard errors, by country and industry.

Keywords: Export Unit Values; Export Quality; Revealed Preference; Vertical Differentiation; Horizontal Differentiation

JEL classification: F1; F2; F4

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1. Introduction

This web-based technical appendix provides additional information related to the results reported in "Estimating Cross-Country Differences in Product Quality". Section 2 discusses issues related to dataset construction. Sections 3, 4 and 5 contain additional first-stage results and compare the first-stage estimates reported in the paper to estimates derived from alternate estimators. Section 5 reports bootstrap estimates of our second-stage estimates. Finally, sections 6 and 7 reports additional second-stage results as well as a comparison of our second-stage results to alternate estimates of quality.

2. Dataset Construction

2.1. Net Trade

We use trade data from the United Nations Commodity Trade Statistics Database (COMTRADE) to construct countries' trade balances with the world (i.e., the other countries in the sample) during the sample period.[1] Trade balances are computed for overall manufacturing, by one-digit SITC manufacturing industry, and for Textiles (two-digit SITC 65) and Apparel (two-digit SITC 84).

Records in the COMTRADE data track reporting countries' bilateral trade flows with each of their partner countries by industry and year. Our approach is to subtract each country's total reported imports from its total reported exports by industry and year.[2] Here, we note caveats and adjustments. Our exact algorithm for refining the trade data can be found by examining the Stata programs used to construct the trade balances.

- *Missing data.* One-year gaps in reporting (i.e. the absence of data for imports, exports or both in a particular country-year) occur in a number of countries that otherwise exhibit regular reporting. All of the reports for Japan and Pakistan,

[1] COMTRADE data are revised over time. The data described here were accessed on June 8, 2006 via the website http://unstats.un.org/unsd/comtrade.

[2] 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.

for example, are missing for 1992 and 1994, respectively. Short gaps in trade reporting were interpolated from the closest years available.

- *CIF versus FOB:* Imports are reported CIF while exports are reported FOB. Exports are adjusted by an estimated year-sector-country-pair transport cost spread.

- *Entrepôt trade.* Hong-Kong and Singapore act as entrepôts for exports from China and Malaysia, respectively. In COMTRADE data, countries importing goods from either China or Malaysia via Hong-Kong or Singapore attribute the entire value of the import to the country of origin, although value is added by the entrepôt. Reports from both exporting countries in turn reflect this addition of value: exports from the country of origin (entrepôt) to the final destination are lower (higher) than reported by the country of destination. To avoid error in origin attribution, we assign preference to reports from Hong-Kong, Singapore, China and Malaysia over those of other countries in exports as well as imports.

- *COMTRADE does not include Taiwanese trade data.* Taiwan is excluded from COMTRADE. We identify Taiwanese trade from flows reported by all countries in the database in which the partner is classified under UN code 490 ("other Asia, not elsewhere specified"), which reporting countries generally use to classify trade with Taiwan. However, given countries' standard practice of reporting imports under the "country-of-origin" criterion noted above, employing partner reports in this manner would lead to double counting of Taiwan's exports to areas in which there is entrepôt trade. A Taiwanese export that passes through customs in Singapore and is later re-exported to Malaysia, for example, would be registered by both countries as an import from UN code 490. To avoid double counting, we employ Taiwan's own data of exports to the People's Republic of China, Hong Kong, Malaysia, Indonesia and Singapore, downloaded from Taiwan's Bureau of Foreign Trade (eweb.trade.gov.tw).

- *Missing reports of Singapore's trade with Indonesia.* Singapore does not report any trade with Indonesia on COMTRADE prior to 2003, even though Indonesia is one of its main trade partners. We employ the ratio of Singapore and Indonesia's trade reports for a same trade flow in 2003 and 2004 to create sector-specific adjustment factors. We proxy Singapore's missing trade reports as Indonesia's factor-adjusted reports up to 2002.

- *Unspecified origin.* Imports reported to originate in general unspecified areas – i.e., "bunkers" (UN code 837), "free zones" (838), "special categories" (839) and "areas not elsewhere specified" (899) – are attributed to partners reporting exports in excess of specified import value. For a country-year-sector in which unidentified imports are greater than the sum of bilateral gaps between specified imports and partner exports, the entire value of unidentified imports is allocated to close the gaps, and any remainder kept as unspecified imports. Where unspecified imports are insufficient to close all gaps, unspecified imports are distributed across partners in proportion to gap value. Exports to unspecified areas are not attributed: because a reporter's fob export value should always be below partners' cif values there is no natural basis for attribution.

- *Non-ferrous metals.* Following standard practice, we remove non-ferrous metals (two-digit SITC 68) from manufactured goods. Products in this category are generally considered commodities. We note that removing trade values at the two-digit level from one-digit data is non-trivial because quality in reporting degrades with disaggregation. We proxy for gaps in reporting because otherwise the construction of bilateral trade flows in one-digit SITC 6 would employ data from different directions of trade when netting imports in two-digit SITC 68. In a number of cases, that would lead to substantial mismatches. A gap in the two-digit record of a given country A with partner B is adjusted if it meets three conditions: i) country A reports a flow from B under one-digit SITC code 6 but not under two-digit code 68, ii) A reports a flow of non-ferrous metals from B in the same direction of trade for some other year during the period, and iii) B claims the existence of the flow in that particular year. We construct an average ratio of flows in two-digit SITC 68 to flows in one-digit SITC 6 for the country pair. To proxy A's report at two-digits, we adjust its one-digit reported value by this average ratio. The average ratio is obtained from the two years surrounding the missing data. If a ratio was unavailable from either surrounding year, the ratio from the other was used. If a ratio in both surrounding years was unavailable, ratios from the closest years available were used.

2.2. *Trade Costs*

Tariff information comes from the Trade Analysis and Information System (TRAINS) Database maintained by the United Nations Conference on Trade and Development (UNCTAD). Using the World Integrated Trade Solution (WITS) internet interface, we download publicly available data on reporting countries' most favored nation (MFN) and bilateral preferential (PRF) tariff rates for pooled manufacturing goods as well as for manufacturing goods by one-digit SITC industry and select two-digit SITC industries (see main text).[3]

Countries assign tariffs according to six- to ten-digit Harmonized System product classifications. Here, we make use of UNCTAD-constructed weighted average tariff rates computed across all tariff lines available for pooled manufacturing and for manufacturing by one-digit SITC industry and for select two-digit SITC industries. These weighted average tariffs use reporting countries' import values as weights.[4]

Unfortunately, the tariff data are available only sparsely. Table 1, for example, reports MFN data availability for pooled manufacturing across the countries in our sample, where EUN refers to the countries of the European Union. Each cell of the table either contains the reporter's weighted average MFN tariff or is missing (i.e., "."). As indicated in the table, MFN data is available for just under half of the cells. Data availability for one-digit manufacturing industries as well as for bilateral preferential tariffs is similarly limited.

Our construct of a trade-cost dataset has three steps. First, we construct a balanced panel of MFN tariffs for both pooled manufacturing and for one-digit manufacturing industries. In both cases, we use the following algorithm. Start with the raw data from TRAINS. Fill in missing reporter-year (MFN) observations by using data for the last year available. For example, if country c reports a pooled-manufacturing MFN tariff of 10 percent in 1995 and 8 percent in 2000, we assume it has pooled-manufacturing MFN tariffs of 10 percent and 8 percent for years 1995 to 1999 and for 2000 to 2004, respectively. Missing observations before the first year of data availability are filled in with the next available observation.

[3]The data can be accessed at http://wits.worldbank.org via Microsoft's Internet Explorer after installing the WITS software (http://wits.worldbank.org/install.htm) and registering with the World Bank. For MFN tariffs, we use reporting countries MFN tariff *vis a vis* the world. For PRF tariffs, we use the preferential tariffs countries report separately for each partner country.

[4]We do not observe changes in the set of tariff lines used to construct these averages across either time or country pairs. As a result, they are likely influenced by composition.

Country	1989	1990	1991	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001	2002	2003
ARG	.	.	.	14	14	.	12	14	15	16	15	15	14	13	14
AUS	.	.	12	.	10	.	.	5	5	5	5	5	4	5	5
AUT	.	11	.	.	.	.	.	.	.	.	.	.	.	.	.
BRA	37	28	23	21	16	16	14	16	16	18	16	15	12	12	12
CAN	8	.	.	.	8	.	7	5	5	4	3	3	3	3	3
CHE	.	0	.	.	0	.	0	0	0	0	0	0	0	0	0
CHL	.	.	.	11	11	11	11	.	11	11	10	9	8	7	.
CHN	.	.	.	36	33	30	.	18	15	15	14	13	13	.	7
COL	.	.	6	10	.	11	12	11	11	.	10	11	10	11	.
EUN	6	7	7	7	7	7	6	5	5	4	3	3	3	4	4
FIN	.	.	.	.	.	.	.	.	.	.	.	.	.	.	.
HKG	.	.	.	.	.	.	.	.	.	0	.	.	.	.	.
HUN	.	.	12	.	10	.	.	9	9	.	.	.	.	8	.
IDN	15	16	.	.	14	.	13	9	.	.	8	7	5	6	6
IND	.	71	.	44	.	.	.	.	21	.	32	.	28	.	.
ISR	.	.	.	.	5	.	.	.	.	.	.	.	.	.	.
JPN	4	3	3	3	3	3	3	3	2	2	2	2	2	2	2
KOR	13	11	.	10	.	.	7	8	.	.	6	.	.	5	.
MAR	.	.	.	.	55	.	.	.	18	.	.	26	26	26	25
MEX	.	.	13	.	.	.	12	.	12	12	15	15	14	15	14
MYS	.	.	11	.	9	.	.	6	6	.	.	.	5	5	5
NOR	.	.	.	.	6	.	6	5	.	2	.	2	2	1	1
NZL	.	.	.	10	9	.	.	8	6	5	5	4	.	5	5
PAK	.	.	.	.	.	.	49	.	.	44	.	.	21	18	17
PHL	23	15	.	15	15	15	14	.	.	6	6	4	3	2	2
POL	.	.	.	11	.	.	10	13	10	10	9	10	8	8	8
ROM	.	.	18	.	.	.	.	.	.	.	14	.	15	.	.
SGP	1	.	.	.	.	.	0	.	.	.	.	.	0	0	0
SWE	5	.	.	.	.	.	.	.	.	.	.	.	.	.	.
THA	35	.	34	.	37	.	16	.	.	.	.	10	11	.	11
TUR	.	.	.	.	8	.	7	.	6	.	5	.	.	.	5
TWN	11	.	.	6	.	.	.	4	.	.	5	3	3	3	3
ZAF	.	11	12	.	14	.	.	9	6	.	5	.	6	.	.

Notes: Table displays mean weighted average MFN tariff by reporting country and year. "." denotes data unavailability. EUN denotes European Union countries.

Table 1: MFN Data Availibility in the TRAINS Dataset, 1989 to 2003

Second, we construct a similar panel of preferential tariffs but omit the final step, i.e., we do not assume preferential tariffs exist before the first year they show up in TRAINS. Our assumption is that preferential tariffs are most likely to appear around the time they are introduced or changed, and casual analysis of the data appears consistent with this assumption.[5]

Third, fill in a complete bilateral database for years 1989 to 2003 using the constructed MFN dataset described above. Then, using the constructed PRF dataset just described, replace the MFN tariff with a PRF tariff if it is available.[6] The final dataset, therefore, contains a mixture of MFN and PRF tariff rates under the assumption that the country pair is governed by MFN rules if preferential rates do not appear in TRAINS. These tariff rates are combined with the estimated transport

[5] The preferential tariffs associated with Mercosur, for example, first show up in 1995.

[6] We keep the MFN tariff in the small number of cases in which the MFN tariff is lower than the PRF tariff.

costs described in the main text to compute the structural trade cost parameters used in the estimation.

Our trade barrier data have three known weaknesses. First, they are available for less than half of the reporter-partner-year cells in our sample period. As a result, we must impute a large number of MFN and PRF tariffs using neighboring values. Second, due to changes in product categorization over time as well as uneven reporting of tariffs across countries and years, the weighted-average tariff rates we rely upon are based upon a non-constant mix of products, introducing potential composition bias. Finally, the TRAINS data do not provide meaningful information about countries' non-tariff barriers, such as apparel and textile quotas.

2.3. *Expenditure shares*

Expenditure shares b_s are defined as the expenditure on sector s as a fraction of total expenditure. Total expenditure is computed, for each country and year, as total GDP minus the total trade balance. Expenditure on sector s is calculated as value added minus net trade in sector s. Information on GDP and total trade balances come from the World Bank's World Development Indicators (WDI) database. Manufacturing value added is computed using data from the United Nations' National Accounts Official Country Data. Value added data at the 3-digit ISIC level are drawn from the United Nation Industrial Development Organization (UNIDO) and converted to SITC sectors using our own concordance largely based on the reverse concordance by Maskus (1989). For each sector s, we use the average expenditure share across countries and years.

2.4. *Real Exchange Rates*

To fill in the missing EIU real exchange rate data discussed in Section 5.2.1, we normalize the World Bank real exchange rate data to the EIU data in overlap years and use the World Bank data for years missing in the EIU dataset. Comparisons of the series in the overlap years indicate a relatively close match and are available upon request. RER data for any remaining holes in the dataset are computed using raw nominal exchange rate and CPI deflator data available in the International Financial Statistics published by the IMF. These data are available at www.imf.org.

3. Additional First-Stage Results

Table 2 reports the share of correctly ordered Paasche-Laspeyres bounds for each sector and year.

	All Manuf	5 - Chemicals	6 - Manuf Mat	7 - Machinery	8 - Misc Manuf
1989	98.6	99.0	98.3	99.0	95.1
1990	99.2	98.8	97.1	98.3	96.6
1991	99.2	99.2	97.4	97.9	96.9
1992	99.1	98.5	96.6	97.4	96.7
1993	98.8	99.4	97.4	97.8	97.4
1994	99.6	99.0	98.1	96.9	97.5
1995	99.1	98.4	97.9	98.4	97.3
1996	99.7	99.5	98.4	99.1	96.9
1997	99.4	99.0	97.7	97.4	97.9
1998	99.0	99.6	97.3	97.5	98.6
1999	98.8	99.2	98.8	96.9	97.6
2000	99.3	98.6	98.7	97.5	98.2
2001	99.6	98.7	99.0	99.1	98.5
2002	99.0	97.0	98.5	98.6	99.0
2003	99.4	94.5	98.3	98.3	98.8
Mean	99.2	98.6	98.0	98.0	97.5

Notes: Table compares the share of correctly-ordered Paasche and Laspeyres bounds, by industry and year.

Table 2: Correctly Ordered Paasche-Laspeyres Bounds

4. Alternate First-Stage Estimators

In this section we describe three alternative first-stage estimators and discuss the sensitivity of results to their use. The first, which we refer to as the "V" estimator, is defined by a quadratic penalty function centered at the midpoint (F_s^{cd}) of each country pair's interval,

$$\Psi^V = \sum_c \sum_{d>c} \left(\ln F_s^{cd} - \ln P_s^{cd} \right)^2 . \tag{1}$$

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

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

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

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

Finally, we investigate the use of an index proposed by Hummels and Klenow (2005),

$$HK_s^{kW} = \prod_{z \in I_s^{kW}} \left(\frac{p_z^k}{p_z^W}\right)^{w_z^k}, \quad k = 1, ..., C \tag{3}$$

where

$$s_z^k = \frac{p_z^k q_z^k}{\sum\limits_{z' \in I_s^{kW}} p_{z'}^k q_{z'}^k}, \; s_z^W = \frac{p_z^W q_z^W}{\sum\limits_{z' \in I_s^{cW}} p_{z'}^W q_{z'}^W}, \; w_z^k = \frac{\frac{s_z^k - s_z^W}{\ln s_z^k - \ln s_z^W}}{\sum\limits_{z' \in I_s^{cW}} \frac{s_{z'}^k - s_{z'}^W}{\ln s_{z'}^k - \ln s_{z'}^W}}. \tag{4}$$

HK_s^{kW} compares country k's prices to those of the world (W) over the set of goods country k has in common with the world $(z \in I_s^{kW})$. The world (W) is an aggregate of all countries – in our case the 43 countries in the sample – except for country k. Therefore, the "world" price of product z, p_z^W, is just total world value divided by total world quantity, q_z^W, omitting country k in the calculation.[7] Though this index has the advantage that it can be computed rather than estimated, it has the disadvantage of treating an aggregation of countries as a single entity without theoretical justification. Finally, a bilateral index for country pair c and d, HK_s^{cd}, is computed as the ratio HK_s^{cW}/HK_s^{dW}.[8]

[7]Since country k is not included to calculate world prices, the set of countries in W varies with k. We do not subindex W by k to simplify notation.

[8]Note that since the HK index is a bilateral index applied to a multilateral purpose, it does not satisfy transitivity, i.e., it cannot be obtained from applying equation (3) directly to countries c and d. Formally, $HK_s^{cd} \neq \prod_{z \in I_s^{cd}} \left(\frac{p_z^c}{p_z^d}\right)^{w_z^{cd}}$.

Goodness of fit, i.e., percent of first-stage Impure Price Index estimates that lie within the Paasche-Laspeyres bounds, for these alternate estimates is generally lower than for our preferred estimator. As reported in Table 3, the "V" performs best among the alternatives, with the "sink" a close second. The performance of the "HK" estimator, on the other hand, is poor: on average, just 43 percent of the bilateral Impure Price Indexes lie within the theoretically mandated Paasche and Laspeyres bounds. Similar differences are manifest in the first-stage estimates: though we find a high cross-country correlation between Impure Price Indexes estimated by our preferred estimator and those estimated by the "V" and the "sink" (the average cross-sectional correlation across years is above 0.99 in both cases), the correlation with the computed "HK" indexes is much lower (an average across years of 0.43).[9] Given the similarity of the preferred, "V" and "sink" estimates, it is not surprising that the second-stage quality estimates to which they give rise are also quite similar.

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

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

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

5. Bootstrapping First-Stage Estimates

We assess the potential influence of first-stage estimation error on our second-stage quality estimates using a bootstrap.[10] This procedure indicates that the potential influence of first-stage estimation error on our estimates of γ is small.

[9] In a few cases, for a given country-year the "sink" estimator yields an indeterminate solution over a compact interval. This indeterminacy occurs for one country per year on average. Choosing alternative points within the interval has negligible effects on the cross-country correlations cited in the text.

[10] Unfortunately, our estimation is not amenable to a Murphy-Topel type correction. Consider estimation of $y_i = \widehat{x}_i\boldsymbol{\beta} + e_i$,where $\widehat{x}_i$ comes from a first stage and is a function of data, $\mathbf{w}_i$, and

For the bootstrap, we use $N = 500$ replications of our first and second stages. For each replication, we first draw country-pairs with replacement to estimate first-stage impure price indexes. Then, for each replication in the second stage, we draw countries with replacement and use the IPIs obtained in the corresponding first-stage replication to estimate the second stage. This procedure yields 500 estimates of γ_s. For all manufacturing, we find a mean and a standard error of -0.254 and 0.091 for γ_s versus -0.241 and 0.084 in Table 3 of the main text. Additional bootstrap results for SITC industries 5, 6, 7 and 8 are reported in Table 4.

	Bootstrap Estimates of γ				
	All Manuf	5 - Chemicals	6 - Manuf Mat	7 - Machinery	8 - Misc Manuf
Mean	-0.254	-0.098	-0.191	-0.099	-0.071
Standard Deviation	0.091	1.904	0.742	0.045	0.050
5th-95th Percentile Interval	-0.421, -0.142	-0.433, 0.285	-0.509, -0.059	-0.191, -0.044	-0.181, -0.02
Iterations	500	500	500	500	500

Notes: Table summarizes bootstrap estimates of γ using the procedure described in the text.

Table 4: Bootstrap Results

Given that sampling with replacement under the bootstrap results in some countries dropping out of the estimation in each iteration, we cannot compute bootstrap standard errors for our quality intercepts and slopes and therefore do not adjust any of the standard errors reported in the main text.

6. Alternate Second-Stage Estimates

To assess the potential sensitivity of our estimates to our use of sectoral global net trade to identify "demand", we examine here an alternate approach which uses countries' exports to the United States. This alternative has two arguments in its favor. First, given that trade data with the U.S. are available at the product level, use of these data would avoid the need to develop the aggregation results that currently occupy a large portion of the paper. Second, U.S. trade data also make it relatively easy to compute "revealed" product-country U.S. import tariffs that are of substantially higher quality than the sector-country MFN tariff information in TRAINS.

The principal argument in favor of the approach used in the paper, on the other hand, is that use of countries' global net trade mitigates the possibility that estimated

estimated parameters, $\widehat{\boldsymbol{\delta}}$, where $\widehat{x}_i = \mathbf{f}(\mathbf{w}_i, \widehat{\boldsymbol{\delta}})$. In our case, $\widehat{x}_i$ is the Impure Price Index estimated for country i. In a standard setting (e.g., Wooldridge 2002, page 139), $\widehat{x}_i$ is estimated based on variables observed at the level of i units (e.g., countries). Here, however, the IPI is estimated based on data at the level of country pairs (i.e., Paasche and Laspeyres bounds). Further, $\widehat{x}_i$ are not a function of data $(\mathbf{w}_i)$ and estimated parameters $(\widehat{\boldsymbol{\delta}})$, but are instead the estimated parameters themselves.

quality trends respond to idiosyncratic country-specific factors such as unobserved trade barriers or liberalizations. Here we provide additional evidence to back up this assertion. First, we develop the theory for how to estimate quality based on countries' exports to the United States. Second, we implement this alternative method empirically and compare the resulting quality estimates it yields to those reported in the paper, identifying the main determinants of the observed discrepancies.

We want to derive U.S. imports from country k in sector z, x_z^k. Based on the assumptions in the paper, U.S. demand is given by:

$$x_z^k = n_z^k q_z^k = n_z^k \frac{\left(\widetilde{p}_z^k \tau_z^k\right)^{1-\sigma_s}}{\left(G_s^{US}\right)^{1-\sigma_s}} b_s E^{US}.$$

Taking logs yields

$$\ln x_z^k = c + \ln n_z^k + (1-\sigma_s)\ln \widetilde{p}_z^k + (1-\sigma_s)\ln \tau_z^k, \tag{5}$$

where $c = \ln\left(\frac{b_s E^{US}}{(G_s^{US})^{1-\sigma_s}}\right)$.

As in the approach used in the paper, this alternative approach requires assumptions about the number of varieties countries produce (n_z^k). Thus, analogous to Assumption 5 in the paper, we here assume that $n_z^k / Y^k = \mu_z \left(\widetilde{p}_z^k\right)^{-\eta_s}$, which implies that $\ln n_z^k = \ln \mu_z + \ln Y^k - \eta_s \ln \widetilde{p}_z^k$.

Substituing this expression into (5) yields

$$\begin{aligned} \ln x_z^k &= c + \ln \mu_z + \ln Y^k + (1-\sigma_s-\eta_s)\ln \widetilde{p}_z^k + (1-\sigma_s)\ln \tau_z^k \\ &= c + \ln \mu_z + \ln Y^k + (1-\sigma_s-\eta_s)\left(\ln p_z^k - \ln \lambda_s^k - \ln \xi_z\right) + (1-\sigma_s)\ln \tau_z^k. \end{aligned}$$

Taking income to the left hand side, we obtain

$$\ln \frac{x_z^k}{Y^k} = c + \varphi_z + (1-\sigma_s-\eta_s)\ln p_z^k - (1-\sigma_s-\eta_s)\ln \lambda_s^k + (1-\sigma_s)\ln \tau_z^k, \tag{6}$$

where $\varphi_z = \ln \mu_z - (1-\sigma_s-\eta_s)\ln \xi_z$ is a product fixed effect (defined at the 10-digit HS level).

We decompose the path of country $k's$ quality relative to the numeraire as $\ln \lambda_{st}^{ko} = \alpha_{s0}^{ko} + \alpha_{s1}^{ko} t + \varepsilon_{st}^{ko}$, which implies

$$\ln \lambda_{st}^k = \alpha_{s0}^{ko} + \alpha_{s1}^{ko} t + \varepsilon_{st}^{ko} + \ln \lambda_{st}^o.$$

We assume that trade costs (τ_z^k) have an observed component, f_z^k, and an unobserved component, θ_z^k, so that $\tau_z^k = f_z^k \theta_z^k$. Substituting into (6) and adding a time index where appropriate, we have

$$\begin{aligned}\ln \frac{x_{zt}^k}{Y_t^k} &= c_t' + \varphi_z + (1-\sigma_s-\eta_s)\ln p_{zt}^k + (1-\sigma_s)\ln f_{zt}^k \\ &\quad -(1-\sigma_s-\eta_s)\,\alpha_{s0}^{ko} - (1-\sigma_s-\eta_s)\,\alpha_{s1}^{ko}t - (1-\sigma_s-\eta_s)\,\varepsilon_{st}^{ko} + (1-\sigma_s)\ln\theta_z^k,\end{aligned}$$

where $c_t' = c_t - (1-\sigma_s-\eta_s)\ln\lambda_{st}^o$. This yields the estimating equation

$$\ln \frac{x_{zt}^k}{Y_t^k} = c_t' + \varphi_z + (1-\sigma_s-\eta_s)\ln p_{zt}^k + (1-\sigma_s)\ln f_{zt}^k + \alpha_{s0}'^{ko} + \alpha_{s1}'^{ko}t + \theta_{zt}'^{ko}, \quad (7)$$

where $\alpha_{s0}'^{ko} = -(1-\sigma_s-\eta_s)\,\alpha_{s0}^{ko}$, $\alpha_{s1}'^{ko} = -(1-\sigma_s-\eta_s)\,\alpha_{s1}^{ko}$, and $\theta_{zt}'^{ko} = -(1-\sigma_s-\eta_s)\,\varepsilon_{st}^{ko} - (\sigma_s-1)\ln\theta_z^k$ is the disturbance term. As in the approach used in the paper, we estimate equation (7) by 2SLS using the real exchange rate as the instrument for the price.

A comparison of second-stage results using country-sector global net trade (top panel; from Tables 3 and 6 in the revised draft) versus results using country-product exports to the United States (bottom panel) is provided in Table 5 below. Although the magnitudes of the estimated coefficients are not directly comparable, both approaches exhibit the expected negative and statistically significant relationship between "demand" and "prices".

Table 5: Estimation Results From Using Countries' Exports to the United States

	Approach Used in Paper				
	0 - All Manuf	5 - Chemicals	6 - Manuf Mat	7 - Machinery	8 - Misc Manuf
Impure Price Index	-0.241 ***	0.089	-0.171 ***	-0.090 ***	-0.055 *
	0.084	0.130	0.085	0.041	0.037
Observations	640	400	608	533	610
	0.90	0.91	0.77	0.89	0.94
R^2	37.7	0.01	4.21	34.3	13.6

	Alternate Approach Using Exports to the United States				
	0 - All Manuf	5 - Chemicals	6 - Manuf Mat	7 - Machinery	8 - Misc Manuf
ln(UV)	-4.390 ***	-4.272 ***	-4.368 ***	-3.108 ***	-5.642 ***
	1.050	1.862	1.052	0.626	1.894
Trade Costs	-0.866 ***	-1.081 ***	-0.846 ***	-0.585 ***	-0.936 ***
	0.246	0.379	0.240	0.089	0.437
Observations	663,270	105,923	209,013	159,921	188,223
Countries	42	40	41	40	42
First-Stage F Statistic	17.7	4.6	26.6	23.6	8.2

The correlations of country quality levels implied by these estimates are positive and statistically significant across countries in all years of the sample. The average of these correlations across years is 0.50 for overall manufacturing, 0.48 for chemicals, 0.56 for manufactured materials, 0.52 for machinery and 0.49 for miscellaneous manufacturing.

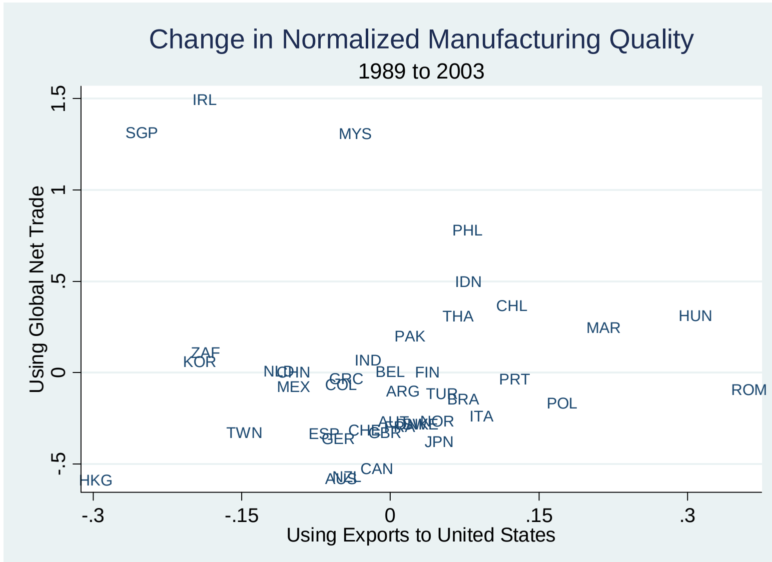

Figure 1: Change in Normalized Quality for Estimates Based on Global Net Trade versus Exports to the U.S.

Comparison of the two methods' implications for countries' change in quality over time is both more interesting and sheds more light on how the two approaches differ. Correlations of countries' 1989 to 2003 quality changes under the two methods, while positive, are not statistically significant. Figure 1 shows the relationship for overall manufacturing. As indicated in the figure, there are notable outliers driving the low the correlation, in particular Singapore, Ireland and Malaysia, for which normalized quality rises substantially more using countries' global net trade than when using countries' exports to the United States as a measure of demand. Romania, on the other hand, is an outlier for the opposite reason.

Intuition for why these countries are outliers (and therefore why the correlation between the two types of estimates is low) is provided in Table 6 below, which compares changes in countries' global net trade to changes in their U.S. exports. In the table, countries are sorted according to the difference between these two measures of changing "demand". As indicated in the first three rows of the table, Singapore, Malaysia and Ireland are the three countries with the largest positive disparity between the two series. Given that these countries' net trade with the world rises more substantially than their U.S. exports, it is not surprising that estimates of quality change based on that measure are larger. Similarly, as indicated in the last row of the table, Romania stands out for having the largest negative disparity between growth in global net trade versus growth of U.S. exports. As a result, estimates of its quality growth over the sample period are stronger when using U.S. exports to identify demand. In sum, discrepancies in estimated quality between the two methods are primarily driven by discrepancies between the two "demand" measures.

Table 6: Change In Global Net Trade versus Change in Exports to the United States, 1989 to 2003

	Change in Net Trade with World Per GDP	Change in Exports to the U.S. Per GDP	Difference		Change in Net Trade with World Per GDP	Change in Exports to the U.S. Per GDP	Difference
SGP	0.274	-0.110	0.384	IND	0.02	0.011	0.009
MYS	0.273	0.133	0.140	ITA	0.009	0.005	0.004
IRL	0.255	0.137	0.118	ESP	0.004	0.001	0.003
IDN	0.114	0.021	0.093	DEU	0.01	0.011	-0.001
THA	0.134	0.046	0.088	MAR	0.002	0.003	-0.001
FIN	0.094	0.011	0.083	TUR	0.004	0.010	-0.006
PHL	0.122	0.064	0.058	JPN	-0.009	-0.002	-0.007
NLD	0.051	0.002	0.049	NZL	-0.003	0.004	-0.007
KOR	0.029	-0.016	0.045	GBR	-0.005	0.002	-0.007
SWE	0.051	0.013	0.038	ISR	0.033	0.047	-0.014
CHE	0.04	0.010	0.030	AUS	-0.014	0.002	-0.016
AUT	0.038	0.009	0.029	ARG	-0.015	0.002	-0.017
CHL	0.026	-0.002	0.028	GRC	-0.019	-0.001	-0.018
CHN	0.077	0.052	0.025	COL	-0.018	0.001	-0.019
TWN	-0.018	-0.042	0.024	BRA	-0.016	0.013	-0.029
BEL	0.029	0.008	0.021	ZAF	-0.03	0.016	-0.046
HKG	-0.058	-0.076	0.018	HUN	-0.03	0.026	-0.056
PAK	0.035	0.018	0.017	CAN	0.003	0.069	-0.066
NOR	0.014	-0.001	0.015	MEX	0.009	0.091	-0.082
FRA	0.018	0.005	0.013	POL	-0.086	0.001	-0.087
DNK	0.017	0.006	0.011	ROM	-0.125	0.008	-0.133
PRT	0.013	0.002	0.011				

Notes: Table displays change in noted series between 1989 and 2003.

We choose countries' global net trade as our preferred measure of demand because we think this measure is less sensitive to unobserved trade costs. Close examination of Canada, which underwent implementation of NAFTA during the sample period, illustrates the advantage of using global net trade rather than exports to the United States. As indicated in Figure 1 above, Canada's normalized quality growth is higher when estimated using exports to the United States than when estimated using global net trade (-0.025 versus -0.337). In Figure 2 below, which complements the information provided in Table 6, we show that the discrepancy in estimated quality trends between the two methods can be closely mapped to a similar discrepancy in the trends of the "demand" variables used to estimate them: While Canada's manufacturing exports to the U.S. (relative to Canada's GDP) rise substantially over the sample period

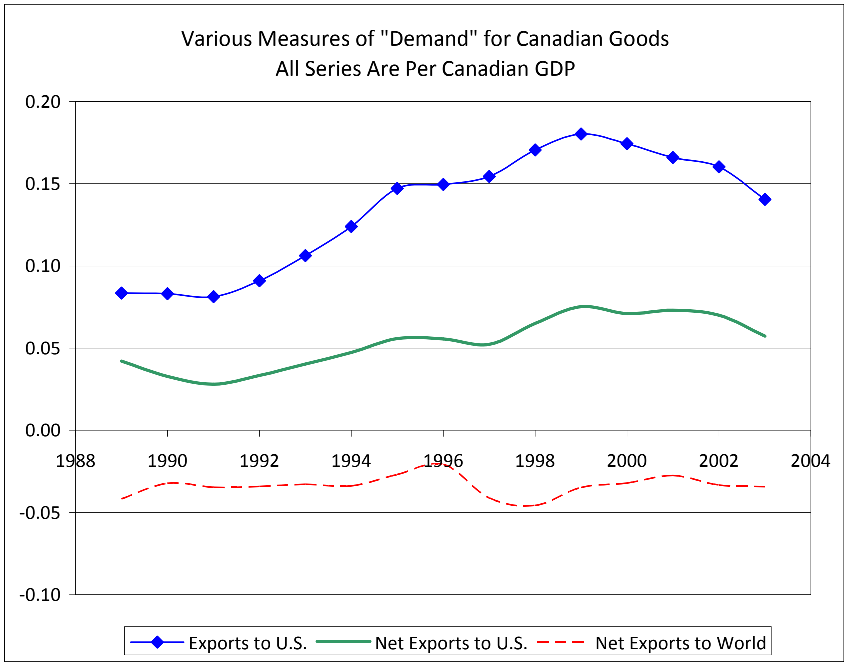

Figure 2: Canada's Net Manufacturing Trade with the World versus Its Exports to the United States

(top line in the figure), Canada's global net trade remains relatively flat. This discrepancy is no doubt due to NAFTA, which substantially increased Canada's exports to the United States but not its net trade with the rest of the world (bottom line in figure). Moreover, in addition to raising Canada's exports to the U.S., NAFTA also raised its imports from this country, with the result that even Canada's net trade with the United States has remained relatively flat (middle line in the figure). Since NAFTA decreased both observable and unobservable bilateral trade costs between Canada and the U.S., using U.S. imports as the "demand" measure mis-attributes to quality the raise in Canada's export to this country that is due to falling unobserved trade costs. In contrast, using global net trade allows for the effect of symmetric unobserved trade costs, or their changes, to cancel out, having no impact on our estimates of quality.

Overall, given the many problems and tradeoffs associated with measuring "demand" empirically, we think it is more justifiable to use global net trade than exports to the United States. First, we believe use of net trade rather than just exports or just imports reduces the potential impact of unobserved changes in symmetric trade costs. Second, use of countries' global trade versus their trade with any particular country (such as the United States) helps mitigate the influence of idiosyncrasies in bilateral trade costs or trade relationships.

7. Additional Second-Stage Results

7.1. Pure Prices and Varieties

Assumption 5 in the paper describes a relationship between export varieties and pure prices:

$$\frac{\overline{n}_s^k}{Y^k} = \left(\widetilde{P}_s^k\right)^{-\eta_s}.$$

This assumption implies that the number of varieties countries export in a particular sector (normalized by GDP) varies inversely with their pure prices in that sector. As a result, variety growth implied by our estimates is the inverse of that implied by our estimates of the Pure Price Index. Malaysia's declining Pure Price Index in Figure 6 of the paper, for example, implies a rise in the number of manufacturing varieties it exports.

Unfortunately, we do not observe the number of varieties countries export within product categories. (For that, we would need to know the number of firms exporting each product from all of the countries included in our analysis, which is not available as far as we know). However, we can observe in U.S. trade data changes in the number of ten-digit HS categories countries export to the United States. While we do not think countries' number of "active" ten-digit HS categories is a sufficiently good proxy for the number of varieties they export to use in our estimation, it can be used to perform a coarse test of Assumption 5.

Toward that end, we estimate a panel regression of country-year normalized export varieties divided by country-year GDP on country-year normalized Pure Price Indexes, with year fixed effects. Results, reported in Table 7, show that our estimated Pure Price Indexes are indeed negatively correlated with observed varieties within years in both manufacturing as a whole and within its constituent sectors. (The relationship between the Pure Price Index and varieties is similarly negative and significant when regressions are performed year by year.)

Table 7: Varieties Exported to the U.S. as a Function of Estimated Pure Prices

	Dependent Variable: Normalized Varieties per GDP			
	0 - All Manuf	6 - Manuf Mat	7 - Machinery	8 - Misc Manuf
Normalized Pure Price Index	-0.853 ***	-0.847 ***	-0.579 ***	-1.079 ***
	0.095	0.146	0.055	0.057
Constant	-26.565 ***	-26.577 ***	-26.633 ***	-26.565 ***
	0.028	0.026	0.033	0.037
Observations	627	595	520	597
Year Fixed Effects	Yes	Yes	Yes	Yes
R-Squared	0.17	0.14	0.21	0.40

Notes: Table reports results of regressing estimated country-year normalized pure price indexes on normalized country-year variety counts per GDP. Pure price indexes and variety counties are in logs; normalization substracts off the log mean of each series in each year, respectively. Varieties are defined as the number of ten-digt HS categories countries export to the United States in the noted sector. *** indicates statistical significance at the 1 percent level.

7.2. Quality Estimates

Tables 8, 9 and 10 report the Quality Index intercepts and slopes for each country in the sample for All Manufacturing, one-digit SITC sectors 5, 6, 7, and 8, and two-digit SITC sectors 65, 84, and 65&84. Intercepts and slopes are obtained from the second stage of the estimation.

	All Manufacturing			
	Quality Intercept		Quality Slope	
Country	Coeff	StdErr	Coeff	StdErr
Argentina (ARG)	-0.113	0.053	-0.007	0.003
Australia (AUS)	0.273	0.018	-0.041	0.007
Austria (AUT)	0.420	0.006	-0.019	0.002
Belgium (BEL)	0.544	0.036	0.000	0.004
Brazil (BRA)	-0.186	0.067	-0.010	0.012
Canada (CAN)	0.209	0.112	-0.037	0.017
Switzerland (CHE)	0.931	0.077	-0.022	0.005
Chile (CHL)	-0.948	0.229	0.026	0.008
China (CHN)	-0.480	0.248	0.000	0.001
Colombia (COL)	-0.454	0.040	-0.005	0.002
Germany (DEU)	0.768	0.340	-0.026	0.014
Denmark (DNK)	0.449	0.035	-0.020	0.002
Spain (ESP)	0.235	0.063	-0.024	0.005
Finland (FIN)	0.666	0.084	0.000	0.004
France (FRA)	0.629	0.053	-0.021	0.000
UK (GBR)	0.473	0.007	-0.023	0.002
Greece (GRC)	-0.414	0.296	-0.002	0.006
Hong Kong (HKG)	0.232	0.255	-0.042	0.013
Hungary (HUN)	-0.157	0.144	0.022	0.002
Indonesia (IDN)	-0.592	0.139	0.036	0.013
India (IND)	-0.523	0.021	0.005	0.002
Ireland (IRL)	0.448	0.019	0.107	0.040
Israel (ISR)	0.377	0.061	-0.026	0.006
Italy (ITA)	0.660	0.143	-0.017	0.006
Japan (JPN)	0.575	0.309	-0.027	0.012
Korea (KOR)	0.168	0.182	0.004	0.003
Morocco (MAR)	-0.604	0.259	0.018	0.011
Mexico (MEX)	-0.329	0.187	-0.005	0.001
Malaysia (MYS)	-0.835	0.006	0.094	0.028
Netherlands (NLD)	0.168	0.035	0.001	0.002
Norway (NOR)	0.180	0.149	-0.019	0.000
New Zealand (NZL)	0.083	0.222	-0.040	0.012
Pakistan (PAK)	-0.690	0.105	0.014	0.002
Philippines (PHL)	-0.740	0.276	0.056	0.022
Poland (POL)	-0.421	0.151	-0.012	0.021
Portugal (PRT)	0.011	0.076	-0.003	0.003
Romania (ROM)	-0.421	0.274	-0.007	0.004
Singapore (SGP)	-0.188	0.068	0.094	0.036
Sweden (SWE)	0.828	0.362	-0.020	0.023
Thailand (THA)	-0.682	0.119	0.022	0.011
Turkey (TUR)	-0.334	0.055	-0.008	0.003
Taiwan (TWN)	0.241	0.260	-0.023	0.002
South Africa (ZAF)	-0.456	0.005	0.008	0.008

Table 8: Quality Index Intercepts and Slopes for All Manufacturing

References

Diewert, W. Erwin and Alice Nakamura, 1993. *Essays in Index Number Theory*, Volume 1. North Holland. Amsterdam.

Feenstra, Robert C., Robert E. Lipsey and Harry P. Bowen, 1997. "World Trade Flows, 1970-1992, with Production and Tariff Data," NBER Working Paper 5910.

Feenstra, Robert C., 2000. "World Trade Flows, 1980-1997," Center for International Data, UC Davis.

	Intercept								Slope				
	5-Chemicals		6-Manuf Mat		7-Machinery		8-Misc Manuf		5-Chemicals		6-Manuf Mat		7-Mac
Country	Coeff	StdErr	Coeff	StdErr	Coeff	StdErr	Coeff	StdErr	Coeff	StdErr	Coeff	StdErr	Coeff
ARG	-0.222	0.223	-0.114	0.041	0.183	0.237	-0.301	0.000	-0.002	0.019	0.002	0.005	-0.012
AUS	0.193	0.035	0.105	0.067	0.476	0.033	0.048	0.485	0.005	0.017	-0.005	0.003	-0.082
AUT	0.143	0.247	0.390	0.145	0.419	0.105	0.210	0.553	0.033	0.003	-0.001	0.004	-0.037
BEL	-0.133	0.335	0.468	0.031	0.848	0.074	-0.155	0.385	-0.013	0.021	-0.008	0.011	-0.049
BRA	-0.358	0.049	-0.105	0.156	-0.154	0.101	-0.509	0.100	0.029	0.043	-0.009	0.014	-0.019
CAN	-0.341	0.137	0.019	0.096	0.383	0.114	-0.207	0.114	0.013	0.047	-0.002	0.005	-0.063
CHE	-0.114	0.521	0.931	0.004	0.903	0.126	0.645	0.074	0.018	0.003	-0.005	0.003	-0.073
CHL	.	.	-0.522	0.184	.	.	-0.818	0.269	.	.	0.021	0.012	.
CHN	-0.247	0.373	-0.614	0.395	-1.289	0.781	0.120	0.183	-0.002	0.005	-0.003	0.001	0.011
COL	.	.	-0.143	0.044	.	.	-0.431	0.020	.	.	-0.005	0.001	.
DEU	0.023	1.387	0.316	0.155	1.216	0.389	0.303	0.495	0.011	0.073	-0.006	0.004	-0.046
DNK	0.161	0.031	0.345	0.003	0.887	0.169	0.319	0.400	0.000	0.011	-0.031	0.016	-0.064
ESP	-0.036	0.033	0.155	0.052	0.318	0.077	-0.012	0.338	0.014	0.009	-0.006	0.006	-0.035
FIN	0.051	0.044	0.500	0.153	0.752	0.090	0.621	0.099	0.044	0.073	0.015	0.004	0.009
FRA	0.084	0.118	0.363	0.050	0.788	0.125	0.520	0.107	0.016	0.024	-0.002	0.001	-0.048
GBR	-0.080	0.227	0.251	0.004	0.726	0.138	0.144	0.450	0.032	0.014	0.003	0.000	-0.056
GRC	.	.	-0.065	0.324	.	.	-0.544	0.716	.	.	-0.014	0.012	.
HKG	-0.055	0.139	-0.662	0.107	-0.857	0.091	3.478	2.799	0.006	0.036	0.004	0.008	-0.050
HUN	.	.	-0.092	0.168	-0.884	0.052	-0.137	0.435	.	.	0.001	0.010	0.105
IDN	.	.	-0.197	0.025	-1.697	0.505	-0.301	0.118	.	.	-0.005	0.000	0.065
IND	0.023	0.425	-0.370	0.003	-0.671	0.100	-0.786	0.175	-0.022	0.037	-0.007	0.002	-0.021
IRL	1.102	1.330	0.485	0.149	0.552	0.219	0.658	0.065	-0.269	0.364	0.003	0.004	0.025
ISR	-0.150	0.819	0.108	0.058	0.845	0.420	0.191	0.139	0.011	0.005	0.014	0.002	-0.057
ITA	0.272	0.337	0.286	0.066	0.481	0.033	0.769	0.259	-0.011	0.013	-0.005	0.003	-0.023
JPN	0.365	1.011	0.280	0.239	0.854	0.570	0.172	0.439	0.018	0.018	0.006	0.005	-0.053
KOR	-0.096	0.694	0.032	0.170	-0.350	0.222	0.569	0.531	-0.003	0.028	-0.006	0.009	0.079
MAR	.	.	.	.	.	.	-0.610	0.558	.	.	.	.	.
MEX	-0.362	0.548	-0.234	0.179	-0.144	0.372	-0.829	0.650	0.021	0.043	-0.017	0.012	0.004
MYS	.	.	-0.472	0.040	-1.699	0.385	0.146	0.747	.	.	0.026	0.018	0.236
NLD	-0.209	0.164	0.102	0.016	0.508	0.336	-0.217	0.183	0.008	0.010	0.002	0.000	-0.037
NOR	-0.019	0.601	0.182	0.035	0.271	0.313	.	.	0.023	0.052	-0.001	0.004	-0.010
NZL	.	.	-0.006	0.480	0.097	0.266	.	.	.	.	-0.016	0.008	-0.077
PAK	.	.	-0.179	0.055	.	.	-0.695	0.195	.	.	-0.009	0.002	.
PHL	.	.	-0.439	0.274	-1.545	0.621	-0.166	0.025	.	.	-0.007	0.012	0.163
POL	.	.	-0.320	0.211	-0.602	0.554	-0.867	0.131	.	.	0.004	0.002	-0.029
PRT	.	.	0.083	0.017	-0.738	0.418	0.668	0.224	.	.	0.001	0.002	0.007
ROM	.	.	.	.	.	.	-0.823	0.841	.	.	.	.	.
SGP	.	.	-0.487	0.013	1.485	1.018	-1.304	0.619	.	.	0.047	0.025	0.100
SWE	0.433	0.357	0.446	0.281	1.481	0.975	-0.060	0.115	-0.010	0.004	0.006	0.008	-0.060
THA	-0.313	0.042	-0.617	0.145	-1.577	0.369	0.416	0.492	0.027	0.053	0.004	0.004	0.071
TUR	.	.	-0.148	0.065	-0.988	0.192	-0.353	0.094	.	.	-0.002	0.006	0.036
TWN	-0.114	0.368	-0.089	0.259	-0.445	0.235	1.342	1.316	0.003	0.023	-0.009	0.001	0.064
ZAF	.	.	0.028	0.050	-0.832	0.500	-1.215	0.395	.	.	0.021	0.009	0.022

Table 9: Quality Index Intercepts and Slopes, by Manufacturing Industry

Hummels, David and Peter Klenow, 2005. The Variety and Quality of a Nation's Exports. American Economic Review, 95: 704-723.

Maskus, Keith, 1989. Comparing International Trade Data and Product and National Characteristics Data for the Analysis of Trade Models, in Peter Hooper and J. David Richardson, *International Economics Transactions.* The University of Chicago Press.

	Intercept						Slope					
	SITC 65		SITC 84		65 & 84		SITC 65		SITC 84		65 & 84	
Country	Coeff	StdErr	Coeff	StdErr	Coeff	StdErr	Coeff	StdErr	Coeff	StdErr	Coeff	StdErr
ARG	.	.	-0.658	0.198	-0.633	0.205	.	.	-0.018	0.005	-0.007	0.003
AUS	-0.360	0.814	0.232	0.534	0.198	0.038	0.016	0.001	-0.013	0.013	-0.013	0.003
AUT	0.516	0.018	0.492	0.208	0.501	0.112	-0.028	0.041	0.018	0.031	0.001	0.004
BEL	0.837	0.153	0.555	0.259	0.431	0.020	-0.026	0.015	0.022	0.020	0.006	0.008
BRA	-0.378	0.064	-0.506	0.042	-0.433	0.046	0.008	0.024	-0.020	0.040	-0.007	0.022
CAN	-0.271	0.024	0.058	0.338	-0.026	0.249	-0.010	0.016	-0.008	0.023	-0.007	0.001
CHE	0.798	0.087	0.750	0.238	0.736	0.130	-0.012	0.023	0.017	0.004	0.003	0.004
CHL	.	.	-0.737	0.190	-0.823	0.231	.	.	0.041	0.026	0.037	0.023
CHN	-0.650	0.695	-0.049	0.558	-0.001	0.260	0.001	0.005	0.005	0.005	0.004	0.000
COL	-0.535	0.188	-0.386	0.065	-0.374	0.050	0.024	0.024	-0.008	0.004	-0.014	0.001
DEU	0.325	0.066	0.832	0.763	0.594	0.575	-0.014	0.019	-0.031	0.040	-0.029	0.031
DNK	.	.	.	.	0.235	0.388	.	.	.	.	0.023	0.021
ESP	-0.028	0.410	0.451	0.077	0.187	0.199	-0.011	0.009	-0.027	0.004	-0.016	0.001
FIN	.	.	.	.	.	.	.	.	.	.	.	.
FRA	0.488	0.069	0.782	0.215	0.731	0.069	-0.013	0.007	-0.011	0.015	-0.016	0.008
GBR	0.265	0.268	0.364	0.243	0.318	0.237	-0.004	0.010	-0.022	0.023	-0.015	0.010
GRC	.	.	-0.338	0.381	-0.412	0.385	.	.	0.023	0.010	0.017	0.005
HKG	-1.649	0.783	2.071	1.572	1.336	1.092	0.089	0.060	-0.133	0.076	-0.076	0.043
HUN	.	.	-0.344	0.278	-0.449	0.191	.	.	0.052	0.002	0.037	0.006
IDN	-0.186	0.058	-0.363	0.173	-0.233	0.102	-0.007	0.002	0.016	0.025	0.023	0.027
IND	-0.510	0.137	-0.661	0.092	-0.567	0.051	0.031	0.048	0.010	0.007	0.021	0.014
IRL	0.055	0.335	0.021	0.424	-0.005	0.403	-0.006	0.003	0.002	0.004	0.001	0.006
ISR	-0.251	0.333	0.125	0.239	0.084	0.201	-0.001	0.007	-0.042	0.024	-0.039	0.022
ITA	0.522	0.064	0.749	0.393	0.779	0.032	-0.003	0.001	-0.006	0.013	-0.007	0.007
JPN	0.113	0.250	0.326	0.453	0.269	0.382	-0.008	0.003	-0.001	0.001	-0.002	0.000
KOR	0.556	0.777	0.311	0.328	0.546	0.457	-0.011	0.003	-0.063	0.028	-0.054	0.022
MAR	.	.	-0.526	0.447	-0.490	0.362	.	.	0.096	0.061	0.039	0.025
MEX	-0.428	0.235	-0.696	0.447	-0.653	0.397	-0.031	0.040	0.003	0.017	-0.003	0.012
MYS	-1.104	0.320	0.133	0.625	-0.036	0.470	0.031	0.033	-0.017	0.011	-0.002	0.001
NLD	0.366	0.301	.	.	0.387	0.169	0.001	0.015	.	.	-0.010	0.013
NOR	.	.	.	.	.	.	.	.	.	.	.	.
NZL	.	.	.	.	0.070	0.495	.	.	.	.	-0.011	0.011
PAK	2.100	1.992	-0.648	0.108	0.382	0.448	0.012	0.036	0.019	0.014	0.026	0.020
PHL	-1.057	0.503	-0.315	0.081	-0.481	0.167	-0.036	0.008	0.022	0.023	0.016	0.019
POL	.	.	-0.820	0.224	-0.846	0.205	.	.	0.059	0.002	0.041	0.012
PRT	-0.274	0.309	0.650	0.005	0.537	0.088	0.026	0.024	-0.018	0.007	-0.014	0.002
ROM	.	.	-1.174	0.915	-1.111	0.823	.	.	0.134	0.085	0.065	0.043
SGP	.	.	-0.130	0.115	-0.603	0.161	.	.	-0.040	0.017	0.002	0.009
SWE	0.256	0.312	0.517	0.571	0.371	0.453	0.021	0.004	-0.030	0.048	0.013	0.018
THA	-0.500	0.153	0.124	0.214	0.102	0.185	-0.017	0.006	-0.012	0.006	-0.011	0.006
TUR	-0.071	0.158	-0.197	0.034	-0.121	0.009	0.000	0.012	0.028	0.023	0.029	0.023
TWN	1.049	1.305	0.068	0.332	0.505	0.540	-0.018	0.015	-0.051	0.033	-0.054	0.031
ZAF	0.005	0.093	-1.064	0.242	-1.004	0.210	-0.002	0.007	0.007	0.012	0.003	0.012

Table 10: Quality Index Intercepts and Slopes for Apparel and Textiles