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6 Bias in U.S. Import Prices and Demand

Robert C. Feenstra and Clinton R. Shiells

6.1 Introduction

Since the work of Houthakker and Magee (1969), it has been known that estimates of the income elasticity of demand for imports to the United States (and to other industrialized countries) are substantially greater than unity. Since these estimates exceed foreign countries' income elasticities of demand for our products, the implication is that balanced world growth will lead to an automatic worsening in the U.S. trade balance. Dissatisfaction with this result has led a number of researchers to suggest that there is an upward bias in the import price indexes and income elasticity estimates, due to the omission of new product varieties or new foreign suppliers of existing products (see Sato 1977; Helkie and Hooper 1988; Hooper 1989; Krugman 1989; and Riedel 1991). According to this argument, over the past several decades the United States has experienced an expansion in the range of imports from rapidly growing, developing countries, but no corresponding decrease in import prices. As a result, the rising share of imports—which is correlated with rising U.S. income—is attributed to a high income elasticity in the import demand equation.

Helkie and Hooper (1988) attempt to correct the estimation of aggregate U.S. import demand by including a measure of foreign countries' capital stocks as a proxy that reflects their movement into new product lines. It would be preferable to incorporate these import varieties from new supplying countries directly into the import price index, and then to estimate the effect on the

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income elasticity. Drawing on the results in Feenstra (1994), we describe in section 6.2 how the appearance of new product varieties, or of new suppliers of existing products, could bias the import price indexes. The major purpose of the paper is to measure this bias over all U.S. imports, and then to determine the effect of this bias on the estimated income elasticity of import demand.

To obtain the import price indexes, the Division of International Prices of the Bureau of Labor Statistics (BLS) surveys importing firms, as described by Alterman (1991). For firms included in these surveys, interviews are conducted to determine the prices of imported goods whose quality characteristics are unchanged over time: we refer to these as "sampled products" and "sampled prices." These interviews necessarily exclude some products from sampled firms and exclude other importing firms entirely. In section 6.2, we argue that if the share of import expenditure on the sampled products is *falling* over time, this will lead to an *upward bias* in the measured index.

The entry of countries into new product lines is one reason to expect that the expenditure on sampled products may be falling, though this can also reflect a more rapid fall in prices from the new suppliers. Both of these hypotheses are consistent with the "product cycle" theory of international trade (Vernon 1966), whereby production of commodities will shift over time to the lowest-cost locations. Thus, the appearance of new suppliers can quite possibly lead to an upward bias in the import price index. This idea is related to the potential bias in the *consumer* price index due to the appearance of new retail outlets offering lower prices (Reinsdorf 1993). Our paper can be viewed as an international analogue to this domestic argument, with new foreign suppliers taking the place of new retail outlets.

In section 6.3, we discuss the sensitivity of our results to three issues: the functional form of the aggregator, the absence of multinational firms, and the availability of firm-level data. While the basic results are derived for a constant elasticity of substitution (CES) aggregator function, we show that similar results can be obtained for the translog case, so the choice of aggregator is not crucial. On the other hand, the results are very sensitive to the assumption that the international transactions being considered are at arm's length, that is, these are not transactions internal to a multinational firm. Since imports internal to the firm are prevalent in some industries, as we describe, the results concerning the bias are not expected to hold in these cases.

The third issue of concern is the availability of data: the correction to the BLS price index described in section 6.2 relies on having data for the expenditure on products sampled from *each importing firm*. This information is not currently collected on a continuous basis. Accordingly, we are forced to rely on country-level rather than on firm-level data. That is, instead of using the expenditure share on sampled products, we will be using the expenditure on *all products* from sampled countries. These import expenditures are obtained from the U.S. Bureau of the Census. Thus, we are relying on the census data to construct proxies for the theoretically correct adjustment to the BLS indexes, which would rely on firm-level data. The usefulness of these proxies will be judged by their statistical significance when included in import demand equations.

In section 6.4, we examine how the adjustments to the import price indexes affect the income elasticity of demand for aggregate U.S. imports. The inclusion of the foreign capital stock proposed by Helkie and Hooper lowers the income elasticity of import demand from about 2.5 to 2.2. In comparison, using the correction based on the falling expenditure share on sampled countries, we find that the income elasticity is reduced from 2.5 to 1.7, or about halfway to unity. Our estimates suggest that the aggregate import price index is upward biased by between 1 and 2 percentage points annually. We conclude our paper by making a simple recommendation on the collection of additional data by the BLS when it interviews firms.

6.2 Potential Bias in the Import Price Index

To motivate our analysis, we consider the case of new retail outlets for domestic goods. Reinsdorf (1993) argues that very similar products will sell at different prices across retail outlets, and cites Denison (1962) to suggest that these price differentials are due to the time lags needed for consumers to respond to the price information, rather than to quality differentials across retail outlets. These new retails outlets are linked into the consumer price index without the price differential being directly incorporated, which results in a potential upward bias in the index. In order to model this bias, it is essential to assume that the similar goods are imperfect substitutes across the retail outlets. This reflects the empirical observation that a lower price at one outlet does not eliminate demand for the same good at another outlet. Reinsdorf and Moulton (chap. 10 in this volume, sec. 10.5) put further structure on the imperfect-substitutes assumption by assuming that the good has a constant elasticity of substitution across the retail outlets.

We will be taking the same approach to modeling the choice of a U.S. firm to import a product from various possible foreign suppliers. That is, we will assume that the U.S. importer treats the different suppliers' products as imperfect substitutes, reflecting any quality differential across suppliers as well as differences in their time lag of delivery, ease of communication, reliability of supply, and so forth. That is, even when observed quality differentials are absent, we will suppose that the wholesale services provided by the various foreign suppliers are enough to differentiate them, from the buyer's point of view. We should stress that the "buyer" in our case is the U.S. *importer* rather than the U.S. *consumer*, since the latter may be entirely unaware of these differences in wholesale services by the various suppliers. We feel that this assumption of imperfect substitution across foreign suppliers is analogous to that made for domestic retail outlets, provided that the import in question is an arms-length transaction between two unrelated firms. In contrast, the import of a product by a multinational from its own production facility abroad would *not* fit into this framework and will have to be treated separately.

6.2.1 CES Index

Like Reinsdorf and Moulton (chap. 10 in this volume, sec. 10.5) we will also assume the buyer treats the product as having a constant elasticity of substitution across the various supplying firms. This assumption is made for tractability, though we will argue in the next section that similar results could be obtained under alternative specifications. With this assumption, the minimum cost of obtaining one unit of services from the foreign suppliers i of some product is given by

$$(1) \qquad c(p_t, I_t) = \left[\sum_{i \in I_t} b_i p_{it}^{(1-\sigma)}\right]^{1/(1-\sigma)}, \sigma > 1,$$

where σ denotes the elasticity of substitution, which we assume exceeds unity; $I_t \subset \{1, \ldots, N\}$ is the set of foreign suppliers in period t with prices $p_{it} > 0$, $i \in I_t$; p_t denotes the corresponding vectors of prices in period t; and $b_i > 0$ denotes a quality (or taste) parameter for the product from supplier i.

Several features of the CES function in equation (1) should be noted. First, we have treated each foreign firm as supplying a single variety i of the differentiated product. Multiproduct firms can be handled, however, by letting i index each variety supplied by each firm. Thus, we will sometimes refer to i as an index of product varieties, where it is understood that this can be across firms or across products within a firm. Second, we have treated the quality parameters b_i as constant over time. This is not essential, and we could alternatively allow these parameters to change. In that case, we would assume that the "quality-adjusted" price is correctly measured for products that the BLS samples: that is, movements in b_i are correctly evaluated for the sampled products. For the nonsampled products, movements in b_i will not affect our results below, because we will use the expenditure shares to evaluate the (unobserved) prices and these shares would also respond to any changes in quality (Feenstra 1994).

To briefly review known results, suppose that the same set of product varieties I are available in periods $t-1$ and t, and that the amounts purchased of each variety, x_{t-1} and x_t, are cost-minimizing quantities for the prices p_{t-1} and p_t, respectively. Let $s_{t-1}(I)$ and $s_t(I)$ denote the corresponding expenditure shares:

$$(2) \qquad s_{it}(I) \equiv p_{it}x_{it} / \sum_{i \in I} p_{it}x_{it}.$$

As in Diewert (1976), the exact price index $P[p_{t-1}, p_t, s_{t-1}(I), s_t(I)]$ is defined as a function of observed prices and expenditure shares, such that

$$(3) \qquad c(p_t, I)/c(p_{t-1}, I) = P[p_{t-1}, p_t, s_{t-1}(I), s_t(I)].$$

The important feature of equation (3) is that the price index itself does not depend on the unknown parameters b_i, $i \in I$. From Sato (1976) and Vartia (1976), a formula for the exact price index corresponding to the CES unit-cost function is

$$(4a) \qquad P[p_{t-1}, p_t, s_{t-1}(I), s_t(I)] \equiv \prod_{i \in I} (p_{it}/p_{it-1})^{w_{it}(I)}.$$

This is a geometric mean of the individual price changes, where the weights $w_{it}(I)$ are computed using the cost shares $s_{it}(I)$ in the two periods, as follows:

$$(4b) \qquad w_{it}(I) \equiv \left[\frac{s_{it}(I) - s_{it-1}(I)}{\ln s_{it}(I) - \ln s_{it-1}(I)} \right] \Big/ \sum_{i \in I} \left[\frac{s_{it}(I) - s_{it-1}(I)}{\ln s_{it}(I) - \ln s_{it-1}(I)} \right].$$

The numerator on the right-hand side of equation (4b) is the logarithmic mean of $s_{it}(I)$ and $s_{it-1}(I)$, and lies between these cost shares. The weights $w_{it}(I)$, then, are a normalized version of the logarithmic means and add up to unity.[1]

The exact price index in equation (4a) requires that the same varieties are available in the two periods, and that the prices for all these products are sampled. We now show how the exact index can be computed when only a subset of the product varieties is sampled. To this end, suppose that I_{t-1} and I_t are the full sets of imported products and that $I \subseteq (I_t \cap I_{t-1})$, $I \neq \emptyset$, is sampled in *both* periods. We shall let $P[p_{t-1}, p_t, s_{t-1}(I), s_t(I)]$ denote the price index in equation (3) that is computed by using data on only this set. We shall refer to this as a "conventional" price index, in the sense that it is computed over a constant set of (sampled) products. The exact price index should equal the ratio $c(p_t, I_t)/c(p_{t-1}, I_{t-1})$. Our first result, proved in Feenstra (1994), shows how this can be measured with observed prices and quantities:

PROPOSITION 1. *For any set of sampled products* $I \subseteq (I_t \cap I_{t-1})$, $I \neq \emptyset$, *the exact price index for the CES aggregator is*

$$(5) \qquad c(p_t, I_t)/c(p_{t-1}, I_{t-1}) \equiv P[p_{t-1}, p_t, s_{t-1}(I), s_t(I)] \times [\lambda(I)_t/\lambda(I)_{t-1}]^{1/(\sigma-1)},$$

$$\text{where } \lambda(I)_r \equiv \sum_{i \in I} p_{ir}x_{ir} \Big/ \sum_{i \in I_r} p_{ir}x_{ir}, \quad \text{for } r = t-1, t.$$

This result states that the *exact* price index equals the conventional index $P[p_{t-1}, p_t, s_{t-1}(I), s_t(I)]$ times an additional term that represents the bias in the conventional index. To interpret this term, note that $\lambda(I)_t$ equals the fraction of expenditure on sampled products in period t, relative to the entire set $i \in I_t$. Thus, $[\lambda(I)_t/\lambda(I)_{t-1}]$ is the ratio of expenditure on sampled products over the two periods. If this ratio is less than unity, reflecting a declining share of expenditure on the sampled products, then the exact price index will be

1. Using L'Hospital's Rule, it is readily shown that as $s_{it-1}(I) \rightarrow s_{it}(I)$ for all i, the weights $w_{it}(I)$ approach $s_{it}(I)$.

lower than the index $P[p_{t-1}, p_t, s_{t-1}(I), s_t(I)]$. In other words, the declining share of expenditure on the sampled products will lead to an upward bias in the conventional index.

A declining share of expenditure on the sampled products could be due to the appearance of new suppliers or, alternatively, to a fall in the relative price of products not included in the sample. Both of these hypotheses are consistent with the product-cycle theory of international trade (Vernon 1966), whereby the production of commodities will shift over time to the lowest-cost locations. Thus, the appearance of new suppliers can quite possibly lead to an upward bias in the import price index. The potential bias in the conventional index is measured by the change in the share of expenditure on the sampled products, raised to the power $1/(\sigma - 1)$. For example, if new suppliers are providing products that are perfect substitutes for existing products, so that σ approaches infinity, then there will be no bias in the existing index. Conversely, if σ is low (but still greater than unity), any given change in the relative expenditure on sampled products will indicate a greater bias in the conventional index.[2]

6.2.2 BLS Index

The BLS samples multiple varieties of a product within each ten-digit Harmonized System (HS) category and then constructs the index at that level. More precisely, given the ratio of prices in the two time periods for each sampled product, the BLS constructs an unweighted arithmetic mean of these prices in the ten-digit HS category; aggregation to broader industry levels then occurs with a Laspeyres formula. The use of an arithmetic rather than a geometric mean will result in some upward bias in the index, and the absence of weights in the index may also introduce some error. In addition to these, we can use proposition 1 to determine the potential upward bias in the BLS index if the sampled products have expenditure shares that are falling over time.

Note that proposition 1 holds even if the set we use to construct the conventional price index P contains only a single variety, so that $I = \{i\}$. In this case the conventional index is simply the price ratio for that single variety, $P = p_{it}/p_{it-1}$, while the term $\lambda_t(i) \equiv s_{it}$ measures the observed expenditure share on that variety. Then taking the geometric mean of equation (5) for all the sampled product varieties $i = 1, \ldots, N$, it follows that the exact price index equals

$$(6) \qquad c(p_t, I_t)/c(p_{t-1}, I_{t-1}) = \prod_{i=1}^{N} (p_{it}/p_{it-1})^{1/N} (s_{it}/s_{it-1})^{1/N(\sigma-1)}.$$

The unweighted arithmetic mean used by the BLS exceeds the simple geometric mean appearing in equation (6). We then obtain

2. The elasticity of substitution must exceed unity, because otherwise all product varieties are essential for consumption, so the set I_t cannot vary over time.

COROLLARY 1. *The BLS index is related to the exact price index by*

(7)
$$\sum_{i=1}^{N} \frac{1}{N}(p_{it}/p_{it-1}) \geq \prod_{i=1}^{N} (p_{it}/p_{it-1})^{1/N} =$$
$$[c(p_t, I_t)/c(p_{t-1}, I_{t-1})] \prod_{i=1}^{N} (s_{it}/s_{it-1})^{-1/N(\sigma-1)}.$$

The final term on the right-hand side of equation (7) is the average decline in the expenditure shares on products sampled by the BLS. When these shares are declining, there is an upward bias in the measured index as compared to the exact index. This bias reflects either the inferred price decline of firms not sampled by the BLS, or the appearance of new product varieties. If we suppose that the newest suppliers—not yet in the BLS sample—also have the most rapidly rising shares, then this upward bias is a plausible outcome. The data used to measure this potential bias are discussed in the next section, after we review the sensitivity of our results to assumptions we have made.

6.3 Sensitivity of Results

6.3.1 Functional Form

The results above were derived under the assumption of a CES aggregator function, and it is important to determine how sensitive the results are to this choice. Suppose instead that the product varieties i enter into a translog aggregator function, so that the unit-cost function in equation (1) is rewritten as

(8)
$$\ln c(p_t, I) = \alpha_0 + \sum_{i \in I} \alpha_i \ln p_{it} + \frac{1}{2} \sum_{i \in I} \sum_{j \in I} \gamma_{ij} \ln p_{it} \ln p_{jt},$$

with $\alpha_i > 0$ and $\gamma_{ij} = \gamma_{ji}$. The set I in this definition refers to the universe of possible product varieties, and is not allowed to vary. For products that are not available in some period, reservation prices, which are generally finite (see below), must be used in the right-hand side of equation (8). This contrasts with the CES case in equation (1), where the reservation prices were infinite, and products that were not available would simply not appear in the unit-cost function. Summing over this universe of products, the unit-cost function is homogeneous of degree one in prices, provided that $\Sigma_i \alpha_i = 1$ and $\Sigma_i \Sigma_j \gamma_{ij} = 0$.

For the translog function, the share of expenditure devoted to variety i is

(9)
$$s_{it} = \alpha_i + \sum_{j \in I} \gamma_{ij} \ln p_{jt}.$$

If there is a variety n that is newly available in period t, then its reservation price in $t-1$ is calculated by setting $s_{nt-1} = 0$ in equation (9), obtaining

(10)
$$\ln \tilde{p}_{nt-1} = \frac{-1}{\gamma_{nn}} \left(\alpha_n + \sum_{\substack{i \in I \\ i \neq n}} \gamma_{ni} \ln p_{it-1} \right).$$

We assume that $\alpha_{nn} < 0$, so that the reservation price is positive and finite for some values of p_{it-1}. This reservation price is used in equations (9) and (10) when variety n is not available.

Our goal is to determine how the translog aggregator would affect the results in proposition 1. To this end, we suppose that variety n is not included in the set of sampled varieties in either period. This may be because variety n is new, or because it is available in both periods but not sampled. In either case, let $I/\{n\} \equiv \{i \mid i \in I \text{ and } i \neq n\}$ denote the set of sampled products. Then the change in the price of variety n between the two periods can be computed from equation (9) as

$$(11) \qquad \ln\left(\frac{p_{nt}}{p_{nt-1}}\right) = \left(\frac{s_{nt} - s_{nt-1}}{\gamma_{nn}}\right) - \sum_{i \in I/\{n\}} \left(\frac{\gamma_{ni}}{\gamma_{nn}}\right) \ln\left(\frac{p_{it}}{p_{it-1}}\right).$$

To interpret equation (11), recall that $\gamma_{nn} < 0$ and that $\Sigma_i \gamma_{ni} = 0$, so that $\Sigma_{i \in I/\{n\}} \gamma_{ni}/\gamma_{nn} = -1$. Then the expression on the right-hand side of equation (11) is a weighted average of the change in prices of all goods $i \neq n$. So equation (11) states that the change in the price of good n, relative to a weighted average of the prices of other varieties, is proportional to the change in the expenditure share on variety n. Note that this expression continues to hold if variety n is not available in one (or both) of the periods, in which case its share is set at zero in equation (11).

To determine the impact of the nonsampled variety on unit costs, we use the result that the ratio of unit costs for the translog function equals a Divisia index of the changes in the individual prices (e.g., Diewert 1976):

$$(12) \qquad \ln[c(p_t, I_t)/c(p_{t-1}, I_{t-1})] \equiv \sum_{i \in I} \frac{1}{2}(s_{it-1} + s_{it})\ln(p_{it}/p_{it-1}).$$

When variety n is newly available in period t, then its reservation price (equation [10]) is used on the right-hand side of equation (12) in period $t-1$. To determine the effect of omitting variety n from the price index in both periods, we substitute equation (11) into equation (12), obtaining

PROPOSITION 2. *Letting $I/\{n\} \equiv \{i \mid i \in I$ and $i \neq n\}$ denote the set of sampled products, the exact price index for the translog aggregator function is*

$$\ln[c(p_t, I)/c(p_{t-1}, I)] =$$
$$\sum_{i \in I/\{n\}} \frac{1}{2}(\tilde{s}_{it-1} + \tilde{s}_{it})\ln(p_{it}/p_{it-1}) - \left(\frac{s_{nt} - s_{nt-1}}{\bar{\eta}_n - 1}\right),$$

where (a) $\bar{\eta}_n \equiv 1 - [2\gamma_{nn}/(s_{nt-1} + s_{nt})]$ is the average elasticity of demand for variety n; and (b) $\tilde{s}_{ir} \equiv s_{ir} - (s_{nr}\gamma_{ni}/\gamma_{nn})$ equals the expenditure share of i if variety n was priced at its reservation level in period r, $r = t-1, t$. If the varieties $i \in I/\{n\}$ are weakly separable from n, then $\tilde{s}_{ir} = s_{ir}(I/\{n\})$ as defined in equation (2).

This result states that the exact index equals the sum of two terms: (i) a Divisia index constructed over the sampled products $i \in I/\{n\}$, where the shares $\tilde{s}_{ir}$ in this index reflect the optimal choice if variety n was not available; and (ii) a term reflecting the change in the expenditure share on variety n and its average elasticity of demand. As a proof, note that from equation (9) the elasticity of demand for variety n in period t is $\eta_{nt} = 1 - (\gamma_{nn}/s_{nt})$. Then the following term appears when equation (11) is substituted into equation (12):

$$\frac{1}{2}(s_{nt-1} + s_{nt})\left(\frac{s_{nt} - s_{nt-1}}{\gamma_{nn}}\right) = -\left(\frac{s_{nt} - s_{nt-1}}{\bar{\eta}_n - 1}\right),$$

where $\bar{\eta}_n \equiv 1 - [2\gamma_{nn}/(s_{nt-1} + s_{nt})]$ is the elasticity of demand computed with the *average* share between periods $t - 1$ and t. This establishes part (a) of proposition 2.

To establish part (b), let p_{nr} denote the observed price for variety n and $\tilde{p}_{nr}$ its reservation price. Holding all other prices fixed, it is immediately clear from equation (9) that $\ln(\tilde{p}_{nr}/p_{nr}) = -s_{nr}/\gamma_{nn}$. Substituting this change in prices into the share equation (9) for s_{ir}, it follows that $\tilde{s}_{ir}$ is the implied expenditure on variety i when n is not available. The shares $\tilde{s}_{ir}$ are not generally observed, which is a limitation of proposition 2. However, if the varieties $i \in I/\{n\}$ are weakly separable from n, then a change in the price of variety n (from its observed to its reservation level), should have no impact on the *relative* expenditure share for varieties $i \in I/\{n\}$. In that case, the formula for the shares in equation (2)—which simply omits variety n from the calculation—would equal $\tilde{s}_{ir}$, so that the Divisia index in proposition 2 can be readily measured.

The condition that the products $i \in I/\{n\}$ are weakly separable from n is rather special, more so because we have already assumed that $\gamma_{nn} < 0$ (so that the reservation price is finite). The latter condition means that the higher-level function defined over the aggregate $i \in I/\{n\}$ and variety n must be translog *but not* Cobb-Douglas. However, this implies that the lower-level function used to aggregate the varieties $i \in I/\{n\}$ *must be* Cobb-Douglas in order for the resulting unit-cost function to be translog.[3] Thus, the varieties $i \in I/\{n\}$ will have constant relative shares. The special nature of this separability assumption is perhaps no worse than the CES case, however, as it is the only function for which every subset of goods is weakly separable from every other. Indeed, it appears to be this separability property, rather than the infinite reservation prices, that makes the analysis of new and nonsampled goods so tractable in the CES case.

In order to compare the translog and CES cases, let us continue to assume that there is a single nonsampled variety n.[4] Then from proposition 1 the exact price index in the CES case is

3. The logic of this statement is that a translog function of translog functions is *not* translog in general: rather, it will involve terms of the form $\ln p_i \ln p_j \ln p_k \ln p_l$, which are ruled out by assuming that either the higher-order aggregator or the lower-order aggregates are Cobb-Douglas.

4. If there are *multiple* nonsampled goods, then we assume that this set of varieties $\{n\}$ is weakly separable from the set $i \in I/\{n\}$, and use the scalar n to denote the aggregate of the nonsampled goods. If these varieties $\{n\}$ originally entered the translog function in equation (8), then the aggregator over them must be Cobb-Douglas, for the reasons discussed in n. 3. Alternatively, we

$$\ln[c(p_t, I_t)/c(p_{t-1}, I_{t-1})] = \ln P[p_{t-1}, p_t, s_{t-1}(I), s_t(I)] + \left(\frac{1}{\sigma - 1}\right) \ln\left(\frac{1 - s_{nt}}{1 - s_{nt-1}}\right) \approx \ln P[p_{t-1}, p_t, s_{t-1}(I), s_t(I)] - \left(\frac{s_{nt} - s_{nt-1}}{\sigma - 1}\right), \tag{13}$$

where the final expression follows if the expenditure share on the nonsampled good is small, so that $\ln(1 - s_{nr}) \approx -s_{nr}$. Comparing equation (13) to proposition 2, we obtain

COROLLARY 2. *If variety n is not sampled, then the ratio of the bias in the conventional index for the translog and CES cases is approximately*

$$\left(\frac{\sigma - 1}{\bar{\eta}_n - 1}\right) \begin{cases} = & 1 \text{ if } \bar{\eta}_n = \sigma, \\ = & 1/2 \text{ if } s_{nt-1} = 0 \text{ and } \eta_{nt} = \sigma, \\ \rightarrow & 0 \text{ as both } s_{nt-1} \text{ and } s_{nt} \rightarrow 0. \end{cases}$$

To interpret the first result above, note that the elasticity of demand for variety n in the CES case is $\sigma(1 - s_{nt}) + s_{nt}$. For small values of s_{nt} this is close to σ, so that if the average elasticity of demand in the translog case equals that in the CES case, then the bias terms are approximately equal. This comparison depends, however, on computing the elasticity of demand $\bar{\eta}_n$ using the average share $(s_{nt-1} + s_{nt})/2$. Alternatively, if variety n is newly available in period t so that $s_{nt-1} = 0$, then the bias term in proposition 2 is written as $s_{nt}/2(\eta_{nt} - 1)$ for $\eta_{nt} = 1 - (\gamma_{nn}/s_{nt})$. With $\eta_{nt} \approx \sigma$, this is about *one-half* the bias in the conventional index $s_{nt}/(\sigma - 1)$ obtained in the CES case. Since these bias terms can also be interpreted as the welfare gain due to the introduction of the new product variety, we have shown that this gain is approximately *twice* as large in the CES case (with $\eta_{nt} \approx \sigma$). Finally, the last result above indicates that these comparisons are quite sensitive to the share of the nonsampled good: if this share approaches zero, then the elasticity of demand $\bar{\eta}_n$ for the translog case approaches infinity and the ratio of the biases approaches zero.

While corollary 2 summarizes the quantitative relation between the biases, an immediate qualitative result from comparing equation (13) and proposition 2 is that for both the CES and translog unit-cost functions, a decrease (increase) in the share of the sampled products indicates an upward (downward) bias in the conventional price index. This result does not rely on the approximation in equation (13), but simply uses the facts that both $\sigma > 1$ and $\bar{\eta}_n > 1$ (since $\gamma_{nn} < 0$). Thus, the qualitative nature of the bias identified in corollary 1—that sampling from firms with a falling expenditure share on their products

could use any aggregator over the nonsampled varieties $\{n\}$ and then just assume that this aggregate enters the translog function in equation (8).

will lead to an upward bias in the index—is preserved across these two functional forms, though the magnitude of the bias will depend on the elasticities of substitution and demand as discussed in corollary 2.

6.3.2 Multinational Firms

An assumption maintained throughout our discussion is that the quantity purchased from foreign firms by the U.S. importer is cost minimizing at the observed prices. This assumption fails to hold, however, when the import is internal to a multinational firm, in which case the transfer price for the import may bear little relation to its economic value. Thus, for these "internal" imports we should not expect the bias we have identified in the conventional index to apply. This conclusion is reinforced by the observation that imports internal to a firm may *not* be differentiated across sources of supply: a U.S. multinational engaged in production abroad at two different plants may very well treat the products from these sources as perfect substitutes. Thus, our other maintained assumption—that imports are differentiated across foreign sources—also fails.

Data on intracompany imports are presented in tables 6.1 and 6.2. In table 6.1, we distinguish U.S. manufacturing imports that are internal to U.S. multinationals (shipped from nonbank U.S. affiliates abroad) from those that are internal to foreign multinationals (shipped to nonbank foreign affiliates in the United States). In addition, we distinguish imports that are intended for sale

Table 6.1 U.S. Imports by Source Companies and Countries, 1982 and 1987 ($ billion)

	1982	1987
Total U.S. nonpetroleum merchandise imports[a]	185.7	366.8
Manufacturing imports from nonbank U.S. affiliates abroad[b]	31.8	57.3
To nonbank U.S. parents	25.4	—
From Canada (transportation equipment)	13.4 (10.6)	—
From Japan	2.2	—
From Mexico	1.6	—
Wholesale trade from nonbank U.S. affiliates abroad[b]	2.7	6.7
To nonbank U.S. parents	2.3	—
Manufacturing imports to nonbank foreign affiliates in the United States[c]	13.8	24.5
From nonbank foreign parent group	—	17.6
From Japan	—	3.9
From Germany	—	3.2
From Canada	—	2.7
Wholesale trade to nonbank foreign affiliates in the United States[c]	58.7	107.3
From nonbank foreign parent group	—	85.1
From Japan (motor vehicles and equipment)	—	53.3 (26.1)
From Germany (motor vehicles and equipment)	—	11.6 (9.2)
From Canada (motor vehicles and equipment)	—	2.5 (—)

Sources: [a]Council of Economic Advisers (1993). [b]U.S. Department of Commerce (1982); Mataloni (1990). [c]Howenstine (1985); U.S. Department of Commerce (1987).
Note: A dash indicates that data were not available.

to consumers (wholesale trade) from those that are intended as inputs into further production (manufacturing imports). The most precise data—dealing with shipments from a company abroad to the same company in the United States—are available from a 1982 or a 1987 benchmark survey.

For U.S. multinationals, the intracompany manufacturing imports amounted to $25.4 billion in 1982, or 14 percent of total nonpetroleum merchandise imports.[5] Of this, $10.6 billion was accounted for by transportation imports from Canada, reflecting the Canada-U.S. auto pact. We have listed the three largest source countries, which were Canada, Japan, and Mexico. There was an additional $2.3 billion of intracompany imports classified as wholesale trade, bringing total intracompany trade from U.S. affiliates abroad to 15 percent of imports. Turning to the foreign multinationals with operations in the United States, the internal manufacturing imports of these firms amounted to $17.6 billion in 1987, or 5 percent of total imports. The three largest source countries are Japan, Germany, and Canada. A much larger amount of imports—$85.1 billion or 23 percent of the total—occurs in wholesale trade.[6] The bulk of this wholesale trade was from Japan, much of which is explained by wholesale trade in automobiles (such as Toyota Motor Corporation sending its vehicles to Toyota Motor Sales, U.S.A.). In total, the intracompany trade of U.S. and foreign affiliates is roughly one-half of total imports.

More detailed evidence for individual industries is provided in table 6.2, which covers only the U.S. affiliates of foreign multinationals and their internal imports in manufacturing.[7] The classification of industries is that used by the Bureau of Economic Analysis (BEA), and the industries are ranked according to the share of internal (i.e., intracompany) imports in total imports. At the top of the ranking are chemicals and primary metals, followed by industrial machinery, household audio equipment, and various food products. The average of the internal manufacturing imports for the entire sample is 8 percent.

The borderline industry in table 6.2 is motor vehicles and equipment, where the internal manufacturing imports are 7 percent of the total. Given the extremely large amount of wholesale internal imports in this industry, we ranked it as *above average* in internal imports, and the same is true for all industries listed above motor vehicles and equipment in table 6.2. Conversely, all industries listed below are treated as *below average* in their internal imports.[8] More specifically, for those industries with an internal-imports share exceeding 8

5. In both 1982 and 1987, imports from majority-owned U.S. affiliates abroad accounted for over 80 percent of the total intracompany imports of U.S. multinationals.

6. Zeile (1993) provides a general description of the merchandise trade of U.S. affiliates of foreign companies, including both manufacturing and wholesale trade.

7. These data are obtained directly from Brainard (1993), whom the authors thank for assistance. Ideally, it would be desirable to have the same data for the internal imports of U.S. multinationals, but this was not as readily available.

8. We judged that tobacco products (which are suppressed in table 6.2) would have above-average internal imports and so they were included in that group.

Table 6.2 **U.S. Affiliates of Foreign Companies: Internal Manufacturing Imports by Industry, 1989**

Classification Code[a]	Internal Imports[b]/Total Imports(%)	BEA Industry Definition
283	46	Drugs
281	28	Industrial chemicals and synthetics
102	22	Copper, lead, zinc, gold, silver
289	20	Chemical products, NEC
353	19	Construction, mining, and materials handling machinery
335	15	Primary metal products, nonferrous
356	15	General industrial machinery
366	14	Household audio, video, and communications equipment
284	12	Soap, cleaners, toilet goods
308	12	Miscellaneous plastics products
349	12	Metal services; ordnance; fabricated metal products, NEC
101	11	Iron ore
265	11	Other paper and allied products
202	10	Dairy products
205	10	Bakery products
208	10	Beverages
291	10	Integrated petroleum refining and extraction
321	10	Glass products
341	10	Metal cans, forgings, stampings
120	9	Coal
209	9	Other food and kindred
305	9	Rubber products
343	8	Heating equipment, plumbing, structural metal products
371	7[c]	Motor vehicles and equipment
355	7	Special industrial machinery
384	7	Medical and ophthalmic instruments and supplies
329	6	Stone, clay, concrete, gypsum, nonmetallic minerals
367	6	Electronic components and accessories
140	5	Nonmetallic minerals, except fuels
220	5	Textile mill products
357	5	Computer and office equipment
379	5	Aircraft, motorcycles, bikes, spacecraft, railroad
381	5	Measuring, scientific, and optical instruments
272	4	Miscellaneous publishing
331	4	Primary metal products, ferrous

(continued)

Table 6.2 (continued)

Classification Code[a]	Internal Imports[b]/Total Imports(%)	BEA Industry Definition
354	4	Metalworking machinery
358	4	Refrigeration and service industry machinery
107	3	Other metallic ores
262	3	Pulp, paper, board mill products
275	3	Commercial printing and services
271	2	Newspapers
342	2	Cutlery, hardware, screw products
352	2	Farm and garden machinery
390	2	Miscellaneous manufacturing
201	1	Meat products
230	1	Apparel and other textile products
250	1	Furniture and fixtures
386	1	Photographic equipment and supplies
010	0	Crops
020	0	Livestock, animal specialties
080	0	Forestry
090	0	Fishing, hunting, trapping
133	0	Crude petrol extraction, natural gas
240	0	Lumber and wood products
287	0	Agricultural chemicals
299	0	Petroleum and coal products, NEC
203	—[d]	Preserved fruits and vegetables
204	—[d]	Grain mill products
210	—[c,d]	Tobacco products
310	—[d]	Leather and leather products
351	—[d]	Engines, turbines
359	—[d]	Industrial and commercial machinery, NEC
363	—[d]	Household appliances
369	—[d]	Electrical machinery, NEC
Average	8	

Source: Brainard (1993).
Note: NEC stands for not elsewhere classified.
[a]Industry code from the Bureau of Economic Analysis.
[b]Includes imports by affiliates only from foreign parent group.
[c]Motor vehicles and tobacco products are treated as having above-average internal sales.
[d] Suppressed by the BEA for confidentiality of individual firms.

percent in table 6.2 (including motor vehicles and equipment), we identified the corresponding three-digit Standard International Trade Classification (SITC) numbers. Excluding petroleum products, there are roughly two hundred three-digit SITC categories, of which about one-half corresponded to those industries listed in table 6.2 with above-average internal imports; the other half are treated as having below-average internal imports. Given this crude division of our sample, our hypothesis is that the bias in the conventional import price index should be more prominent for the industries with below-average internal imports.

6.3.3 Availability of Data

The potential bias in the BLS import price index is measured by the last term appearing in corollary 1, that is, the change in expenditure shares on sampled products. An immediate difficulty with implementing this formula is that the expenditure shares on the sampled products are not collected on a continuous basis by the BLS. While expenditure information is used to form an initial sample, once a product has been selected for a price interview, the firm is no longer asked to report the expenditure on that product. For this reason, we have relied on certain proxies for this bias term, constructed from disaggregate import data available from the U.S. Bureau of the Census, over the period 1978–88. The census import data is reported according to the Tariff Schedule of the United States (TSUSA) classification, which includes over ten thousand categories annually. The extremely disaggregate nature of this data set makes it a useful source for constructing expenditure shares on imports.

We will consider two proxies for the bias term in corollary 1. The first replaces the *firm-level* expenditure shares with the corresponding *country-level* expenditure shares in the same product category. That is, for each three-digit SITC industry, we obtained from BLS a list of the countries from which price data was actually collected. This information was obtained for the interviews conducted at two times—September 1982 and March 1985. We also need to make some assumption about what interviews occurred in other years. In the absence of other information, we will assume that the country-product interviews used in 1982 remained constant over the period 1978–83, and that the country-product interviews used in 1985 remained constant over the (overlapping) period 1983–88.

To describe the first proxy, suppose that the BLS obtained information on product i imported from country $k(i)$, in years $t - 1$ and t. We have used s_{it} in corollary 1 to denote the share of expenditure on product i, relative to all imports in that product category.[9] We only have information on the countries sampled at the three-digit SITC level, so we construct the bias at that level. Letting $s_{k(i)}$ denote the import share of country $k(i)$ at the three-digit SITC level, our first proxy for the bias term appearing in corollary 1 is

$$(14) \qquad \text{SHARE1}_t = \prod_{i=1}^{N} [s_{k(i)t}/s_{k(i)t-1}]^{1/N} = \prod_{k=1}^{K} (s_{kt}/s_{kt-1})^{\omega_k},$$

where this term is constructed for each three-digit SITC industry.

To obtain equation (14), we simply replace the product share s_{it} in corollary 1 with the country shares $s_{k(i)t}$. We have also omitted the elasticity term

9. BLS will sample multiple products within each ten-digit HS category (which have replaced the TSUSA classifications since 1989), so in principle, s_{it} denotes the share within this category.

$1/(\sigma - 1)$ which appears as a power on the bias in corollary 1, since this will be estimated when we include equation (14) as a variable in an import demand equation (as described in the next section). Note that the share of country k is repeated each time an import product i (within the same three-digit SITC category) is interviewed from that country. Then, letting ω_k denote the share of interviews within each three-digit SITC for products coming from country k (which was provided to us by BLS), the second equality in equation (14) is obtained.

Our second measure of the potential bias is closely related to the first, but uses information on the detailed TSUSA-level products supplied by each country. In particular, a country that supplies in more TSUSA categories over time can be judged to have increasing product variety in its exports to the United States. The expected impact of greater product variety would be to reduce the expenditure share s_{it} on each variety supplied by individual firms. In the absence of firm-level data, we can evaluate these changes in product variety by computing the country share $s_{k(i)}$ over *only* those TSUSA categories that country k supplies continuously. That is, for each three-digit SITC category and for each source country, we identified the TSUSA products supplied *every year* in the subperiods 1978–83 and 1983–88. Then we calculated the expenditure on these TSUSA products relative to all U.S. imports in the same three-digit SITC industry: this expenditure share is denoted by s^*_{kt}, which is less than the country share s_{kt} by construction. Greater product variety from country k will mean that s^*_{kt} falls relative to s_{kt}. Our second measure of the potential bias is then

$$\text{SHARE2}_t = \prod_{k=1}^{K} (s^*_{kt}/s^*_{kt-1})^{\omega_k}, \tag{15}$$

where $s^*_{kt} \le s_{kt}$ denotes the expenditure on TSUSA products that country k supplies continuously over 1978–83 or 1983–88, relative to total U.S. imports in the same three-digit SITC category.

We expect that SHARE2 would be a better measure of the potential bias than SHARE1, because it takes into account changes in product variety from each country. A limitation of SHARE2 occurs, however, when the names of the TSUSA categories change over time, as they do in response to product innovations or changes in U.S. trade laws.[10] For example, as televisions of increased variety were imported into the United States, the TSUSA categories adjusted to reflect this (distinguishing color versus black and white, and different sizes of screen). If a TSUSA category is split during our sample period, then we count that product as not continuously supplied and ignore it in the calculation of s^*_{kt}. In principle, our calculation is robust to these changes in TSUSA names: if a product with a fixed percentage of country k's export sales (within some three-digit SITC industry) is omitted from the calculation of s^*_{kt}

10. The TSUSA numbers change very frequently, and for this reason, we ignore the numbers and use only the TSUSA names.

and s^*_{kt-1} because its TSUSA category split, this would have no impact on the ratio (s^*_{kt}/s^*_{kt-1}). However, when many of these changes in product names occur, then this ratio is calculated over a very small number of (continuously supplied) TSUSA products.[11] In that case, we might expect SHARE2 to display more erratic behavior than SHARE1. In general, we will judge the usefulness of these two proxies by their significance in regressions of import demand, as described in the next section.

6.4 U.S. Import Demand

We will follow Helkie and Hooper (1988) in specifying a log-linear equation for aggregate U.S. imports:

$$\ln Q_{mt} = \beta_0 + \beta_1 \ln P_{mt} + \beta_2 \ln P_{dt} + \beta_3 \ln Y_t + \varepsilon_t, \tag{16}$$

where Q_{mt} is real nonpetroleum imports, P_{mt} is the aggregate import price index (based on the BLS interviews), P_{dt} is the U.S. gross national product (GNP) deflator, and Y_t is nominal GNP. Since demand should be homogeneous of degree zero in prices and income, we can impose the constraint $(\beta_1 + \beta_2 + \beta_3) = 0$ on equation (16) and rewrite it as

$$\ln Q_{mt} = \beta_0 + \beta_1 \ln(P_{mt}/P_{dt}) + \beta_3 \ln(Y_t/P_{dt}) + \varepsilon_t, \tag{17}$$

which is the form usually estimated.

In the first row of table 6.3, we show the results of estimated equation (17) with quarterly data over the period 1979:1–1988:4. In addition to the variables in equation (17), Helkie and Hooper include a measure of capacity utilization (in the United States relative to that abroad). The coefficients of the relative import price follow a second-order polynomial with eight quarterly lags, real GNP includes one quarterly lag, and the equation is estimated with first-order autocorrelation. The long-run income elasticity is estimated at 2.5.[12] Helkie and Hooper use an average of foreign countries' capital stock (relative to the U.S. capital stock) as a determinant of their ability to move into new product lines. In the second regression in table 6.3, this relative foreign capital stock lowers the income elasticity to 2.15, though the coefficient of the capital stock is insignificant. Over the longer period 1969:1–1984:4 (used by Helkie and Hooper) this variable is more precisely estimated, though the income elasticity is nearly identical to that in table 6.3.

As an alternative to the capital-stock variable, we will use the bias terms

11. In an extreme case, there might be no TSUSA category within a three-digit SITC in which an interviewed country supplied continuously. When this happened (which was infrequently) we replaced the value of (s^*_{kt}/s^*_{kt-1}) for country k with (s_{kt}/s_{kt-1}) before computing equation (15).

12. If a (linear) time trend is introduced in this equation, its coefficient is 0.002, which is highly insignificant and reduces the income elasticity to 2.25. In contrast, for disaggregate import demand equations, Alterman (1993) argues that the inclusion of a time trend can significantly reduce the income elasticities.

Table 6.3 U.S. Import Demand

Relative Import Price	Real GNP	Relative Capacity Utilization	Relative Foreign Capital	SHAREA	SHAREB	ρ	$\bar{R}^2$
−1.147	2.491	−.030	—	—	—	.535	.993
(0.205)	(0.281)	(.175)				(.143)	
−0.979	2.154	−.157	−1.483	—	—	.476	.994
(0.216)	(0.332)	(.186)	(0.942)			(.151)	
−1.231	1.894	−.016	—	.662	−1.450	.312	.994
(0.175)	(0.475)	(.157)		(.204)	(0.795)	(.165)	
−1.149	1.733	−.105	—	.478	−0.926	.429	.991
(0.226)	(0.953)	(.284)		(.288)	(0.831)	(.169)	

Notes: Standard errors are in parentheses. The dependent variable is the log of the important quantity.

The sample range is 1978:1 to 1988:4. The coefficients of the relative import price follow a second-order polynomial with eight quarterly lags; real GNP includes one quarterly lag; and the relative foreign capital stock is entered as a lagged value.

The third regression uses SHARE1A and SHARE1B, while the fourth regression uses SHARE2A and SHARE2B; in both cases the instruments for this variable are t, t^2, t^3, and the other variables in the regression. Since the share variables are measured annually, quarterly dummies are included as instruments and are also included in the third and fourth regressions above (but not reported).

SHARE1 and SHARE2. We suppose that the correct price to include in the import demand equation (17) is the exact index, which is related to the conventional index by corollary 1.[13] Substituting this into equation (17), we obtain

$$\ln Q_{mt} = \beta_0 + \beta_1 \ln\left(\frac{P_{mt}}{P_{dt}}\right) + \left(\frac{\beta_1}{\sigma - 1}\right)\ln(\text{SHARE1}_t) + \beta_3 \ln\left(\frac{Y_t}{P_{dt}}\right) + \varepsilon_t, \tag{18}$$

where SHARE2 is alternatively used. We take a weighted geometric mean over these variables at the three-digit SITC level to arrive at the aggregate value for SHARE1 or SHARE2, where we distinguish those industries with above-average and below-average intracompany imports (using table 6.2).[14] Thus, SHARE1A denotes the mean of SHARE1 over the industries with above-average imports, SHARE1B denotes the mean over the industries with below-average imports; for SHARE2A and SHARE2B are defined likewise. Using

13. It can be questioned whether using the exact price index in equation (17) also means that the exact quantity index should be used on the left-hand side. We will follow the usual practice of using the real imports obtained by deflating nominal imports by the BLS index, rather than deflating by an exact price index. Note that the issue of how to construct the quantity variable goes away if the share of imports in total expenditure is used on the left-hand side, as in Feenstra (1994), for example.

14. The weights in this geometric mean are the average export values in each three-digit SITC industry over the 1978–83 period or over the 1983–88 period.

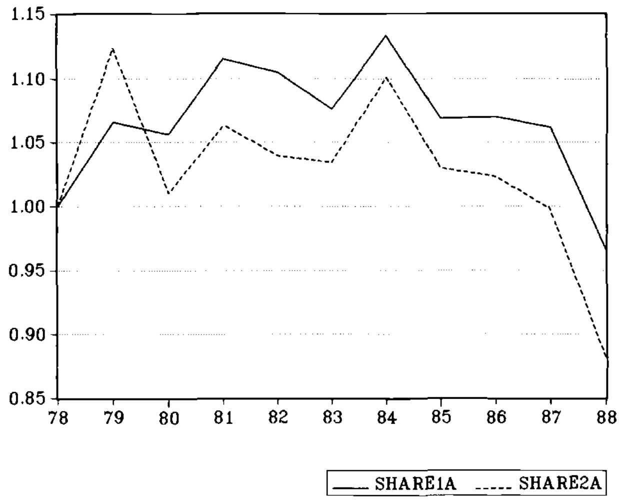

Fig. 6.1 Industries with above-average intracompany imports (shares, 1978–1988)

the aggregates for both groups of industries in equation (18), we arrive at the estimating equation

$$(19) \qquad \ln Q_{mt} = \beta_0 + \beta_1 \ln\left(\frac{P_{mt}}{P_{dt}}\right) + \alpha_1 \ln(\text{SHARE1A}_t) + \left(\frac{\beta_1}{\sigma - 1}\right) \ln(\text{SHARE1B}_t) + \beta_3 \ln\left(\frac{Y_t}{P_{dt}}\right) + \varepsilon_t,$$

where SHARE2A and SHARE2B are alternatively used.

In figure 6.1 we show the values for SHARE1A and SHARE2A, aggregated over industries with above-average intracompany imports, and in figure 6.2 we show SHARE1B and SHARE2B, for industries with below-average internal imports.[15] All the SHARE variables are normalized at 1.0 in 1978. In figure 6.1, the SHAREA variables are quite erratic, showing little trend aside from a decline in the last years of the sample. In figure 6.2, by contrast, the SHAREB variables for industries with below-average internal imports show a marked tendency to decline. SHARE1B reflects the import shares of countries with sampled products, and it declines to 0.88, or about 1 percent annually. A greater decline—to 0.75—is shown by SHARE2B, or about 2.5 percent annually. This fall indicates that the countries with sampled products were also moving into new product lines, so that the expenditure share on the products supplied continuously declined more rapidly.

15. The data for these aggregates are reported in the appendix, table 6A.1.

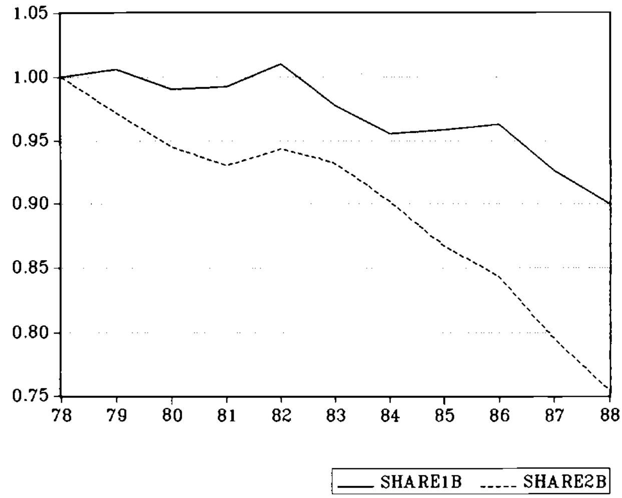

Fig. 6.2 Industries with below-average intracompany imports (shares, 1978–1988)

The results of including the SHARE variables in the import demand equations are reported in the third and fourth regressions in table 6.3, where the third uses SHARE1A and SHARE1B, while the fourth uses SHARE2A and SHARE2B. In both cases, we see that SHAREA enters with a positive sign and SHAREB with a negative sign. The sign on SHAREB is expected, since $\beta_1 < 0$ in equation (19) in the price elasticity of demand, so with $\sigma > 1$ the coefficient on SHAREB is negative. We have not offered any prediction about the sign on SHAREA, however.

One rationalization for the positive coefficient on SHAREA is that when a company decides to shift production offshore, rather than produce domestically, we will observe an increase in both quantity and share of imports from that foreign-country source. Conversely, when a foreign company decides to expand its U.S. manufacturing base, rather than to import, there will be a decline in both the quantity and the share of expenditure from that source country. It is entirely possible that the products internally imported by these companies are included in the BLS interviews, so that the positive correlation between SHAREA and imports is to be expected.[16]

This argument concerning the sign of SHAREA highlights the fact that all the SHARE variables are likely to be correlated with the error in equation (19), since any random change in the import quantity from the sampled countries will also affect their expenditure shares. To address this, the third and fourth regressions in table 6.3 use instrumental variables when including the SHARE

16. A product is excluded from the BLS interviews only if the company states that the import price for that product is not influenced by the market, which seldom occurs.

variables: the instruments are time, time^2, time^3, and the other variables on the right-hand side of equation (19). Since the SHARE variables are measured as annual values, quarterly dummies are also included in the instruments and the regression.

In the third regression in table 6.3, using SHARE1A and SHARE1B, the income elasticity falls from 2.5 to 1.9, and the coefficients of both SHARE variables are significant at the 10 percent level. The autocorrelation coefficient is also reduced. A slightly larger impact on the income elasticity is obtained when using SHARE2A and SHARE2B, calculated according to equation (15). In the fourth regression, the income elasticity falls to 1.7, though the standard errors of the SHARE coefficients are higher than before. The reduction in the income elasticity in either case is the principal result of our paper: the SHARE variables have a substantial effect on the income elasticity of aggregate import demand, moving it about halfway toward unity. This result supports the hypothesis that the high income elasticity of import demand is due, at least in part, to the inability of conventional indexes to account for the expansion of product varieties from new foreign suppliers.

Using the coefficient of SHARE2B in the fourth regression, along with the long-run price elasticity β_1, we can obtain an estimate of σ from equation (19) as $\hat{\sigma} = 1 + (1.149/0.926) = 2.24$ (with a standard error of 1.08). This estimate seems low for an elasticity of substitution between a product differentiated across suppliers, and it is smaller than the disaggregate estimates in Feenstra (1994). One reason for this might be that the SHARE variables are *proxies* for the true expenditure shares from interviewed firms, which could bias the elasticity estimate. For example, if SHARE2B measures only a fraction of the true expansion in product varieties, then this variable would fall too slowly, and the resulting elasticity estimate of $\beta_1/(\sigma - 1)$ in equation (19) will be upward biased—so $\hat{\sigma}$ will be downward biased. More generally, from our results in section 6.3.1, we need not assume that the true aggregator is CES, so that the coefficient of SHARE2B is open to interpretation.

Regardless of how we interpret the coefficients of the SHARE variables, we can combine these two terms with the relative import price and rewrite (19) as

$$\ln Q_{mt} = \beta_0 + \beta_1 \left[\ln\left(\frac{P_{mt}}{P_{dt}}\right) + \left(\frac{\alpha_1}{\beta_1}\right) \ln(\mathrm{SHARE2A}_t) + \left(\frac{1}{\sigma - 1}\right) \ln(\mathrm{SHARE2B}_t) \right] + \beta_3 \ln\left(\frac{Y_t}{P_{dt}}\right) + \varepsilon_t .$$

The term in brackets is our estimate of the (relative) exact import price index. Then using the estimates from the fourth regression in table 6.3, we construct

$$\ln(\mathrm{PRICEB}_t) = \ln(P_{mt}/P_{dt}) + \left(\frac{0.926}{1.149}\right) \ln(\mathrm{SHARE2B}_t),$$

and

$$\ln(\text{PRICEAB}_t) = \ln(P_{mt}/P_{dt}) + \left(\frac{0.926}{1.149}\right)\ln(\text{SHARE2B}_t) - \left(\frac{0.478}{1.149}\right)\ln(\text{SHARE2A}_t).$$

The first of these series only takes account of the industries with below-average intracompany imports, while the second series takes into account all industries. Also, let $\text{PRICE}_t = (P_{mt}/P_{dt})$ denote the (relative) BLS import price index.

In figure 6.3, we plot PRICE, PRICEB, and PRICEAB (with 1978:1 = 100). The fall in PRICE over the period 1980–85 reflects the appreciation of the dollar. Both of the other series lie below PRICE, indicating the upward bias of the conventional index, with PRICEAB lying below PRICEB in all years except 1987–88. The difference between PRICE and PRICEB in 1988 is 16.4, relative to their initial values of 100, while the difference between PRICE and PRICEAB in 1988 is 12.9. Since these differences develop over the decade 1978–88, we conclude that the conventional price index is upward biased by about *1.5 percentage points annually,* as compared to an exact index.

6.5 Conclusions

As a necessary result of the sampling procedure used by BLS to construct (domestic or international) price indexes, some products will be excluded from these indexes. In this paper, we have discussed the consequences of this exclusion. Our basic result is that the expenditure shares on the sampled products provided very useful information on the movement in prices of the nonsampled

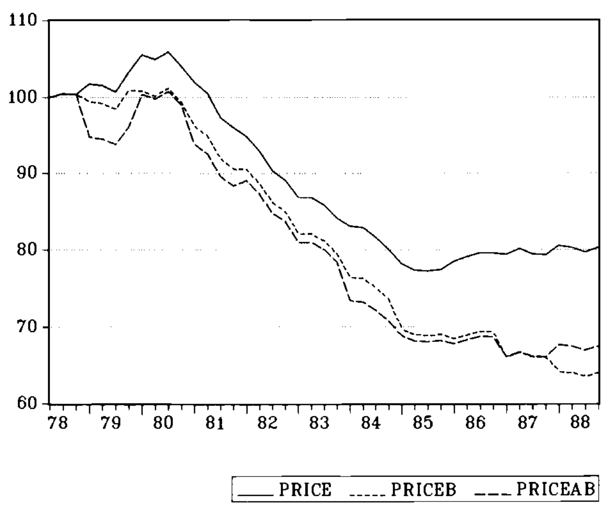

Fig. 6.3 U.S. relative import price, 1978–1988 (1978:1 = 100)

goods. In particular, a falling expenditure share of the sampled products means that we infer a falling relative price for the nonsampled products. This inference is particularly useful when we consider that some of the nonsampled products may be new, with prices falling from their reservation to observed levels when they are first available. Since these reservation prices are never observed (and difficult to estimate when dealing with many goods simultaneously), the strategy of using the expenditure shares to infer the movements in prices seems quite attractive.

In figure 6.3, we have plotted the (relative) U.S. import price index along with two constructed indexes, to illustrate the upward bias in the former. It should be stressed that this diagram is *not* meant to demonstrate any limitation of the BLS procedures in collecting the import price data. Even with the best practice techniques, we would expect any price index constructed from interview data to be potentially biased from the exclusion of products. It would be futile (and prohibitively expensive) to attempt to collect a range of prices broad enough for this potential bias to be eliminated, since the (reservation) prices for new product varieties are simply not available.

Rather than expanding the scope of the price interviews, the recommendation of this paper is that the BLS collect *expenditure* data from firms at the same time as *price* data. Currently, the expenditure on sampled products is not collected on a continuous basis. While expenditure information is used to form an initial sample, once a product has been selected for a price interview, the firm is no longer asked to report the sales (for domestic price indexes) or purchases (for import price indexes) of that product. The collection of this information would impose some extra time costs on the reporting firms, but it would not require any new procedures for selecting the products to interview. That is, once a narrowly defined product has been identified for which to obtain price data, the firm could be asked to supply (quarterly or annual) value data on exactly that same product. These data could be reported at the same level of aggregation as the price indexes, so that the confidentiality of firms is maintained. We have argued that this expenditure data would be very useful for dealing with the potential bias in import prices, and it would undoubtedly be useful for domestic indexes, as well.

Appendix

Table 6A.1 Values of SHARE1 and SHARE2 for Aggregate U.S. Imports

	Industries with Above-Average Internal Imports[a]		Industries with Below-Average Internal Imports[b]	
Year	SHARE1A	SHARE2A	SHARE1B	SHARE2B
1978	1.0000	1.0000	1.0000	1.0000
1979	1.0658	1.1241	1.0060	0.9718
1980	1.0559	1.0098	0.9905	0.9451
1981	1.1155	1.0627	0.9923	0.9303
1982	1.1048	1.0393	1.0102	0.9437
1983	1.0760	1.0342	0.9779	0.9321
1984	1.1334	1.1012	0.9556	0.9018
1985	1.0684	1.0297	0.9586	0.8668
1986	1.0697	1.0230	0.9630	0.8429
1987	1.0615	0.9976	0.9261	0.7952
1988	0.9666	0.8803	0.8999	0.7538

[a]These industries have internal imports greater than 8 percent in table 6.2, including motor vehicle equipment and tobacco products and excluding petroleum products.

[b]These industries have internal imports less than 8 percent or suppressed in table 6.2, excluding motor vehicle equipment, tobacco products, and petroleum products.

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Comment Zvi Griliches

This is an interesting and ambitious paper. It tries to use "share" data on priced commodities to infer the bias that arises from unpriced items. It also argues that import income elasticities are overestimated because of this omission.

There are two parts to the paper. First, the theoretical discussion shows that there is information in the movement of shares of new products about their unobserved "true" prices, *provided* we have or can estimate the relevant price elasticities. In the second part, they try to do just that, estimating the implied

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substitution elasticities by including share data in an import demand equation to adjust for such a bias.

I have a number of comments about this methodology and the empirical results. I am especially interested in the former, since I have been trying to do something similar in my work on generic drug prices (see Griliches and Cockburn 1994). So I will discuss first the "generic" index problem tackled here, then complain about the particular functional form chosen, and then make a few comments on the empirical implementation.

There are two related topics, slightly confused in this literature: "missing prices" and "new distinct goods." The original hedonics literature arose out of the problem that the price of a new good was not available in the base period and looked for a method of retro- and inter-polation. In the spirit of what statistical agencies were doing, it tried to predict what the "market price" would have been for this product yesterday, had it been available, without asking the question whether this was a "demand" or a "supply" price. Once one asks that question, it is clear that the answers can differ, and that integrating the difference between them would yield a measure of consumer's surplus. The theory of exact price indexes and expenditure functions was a tool, developed later, to provide standard index numbers with such an interpretation.

Most of the earlier discussion of new goods was about the appearance of new varieties which were considered to be close substitutes for the previous items: more horsepower, higher speed, and so forth. When the qualitative-choice literature developed, it was natural to recast the problem as one of a distinct new good (choice) and this led to an explicit appearance of a discontinuity in consumer's surplus.

The basic insight used in Feenstra and Shiells's paper is the fact that new goods getting a significant product share implies that in some sense their "real" quality-adjusted price is lower. (This idea is also used in Trajtenberg 1990 and Berry 1994.) Feenstra and Shiells show that the "real" price of the new item must have fallen by

$$dp/p_n = s_n/(\eta - 1).$$

If we can observe the share change, "all" we need is an estimate of the relevant elasticity to find out by how much the price really changed.

So then we are led to the estimation of η (by Hausman, chap. 5 in this volume) or of σ by Feenstra and Shiells. I will come back to that.

The first part of Feenstra and Shiells's paper, as in Feenstra's earlier work, and the conclusions at the end are based on the assumption that the utility structure for varieties is CES. This is convenient computationally, but problematic in practice. First, theoretically, all versions have to be equally substitutable within a nest, and all nests have to be embedded in a Cobb-Douglas function, if there are other CES components. This issue was discussed by Hanoch (1971) twenty years ago in the production-function literature.

Both I (in the paper with Iain Cockburn) and Jerry Hausman, show that the CES calculations lead to implausibly large numbers. Feenstra and Shiells themselves show that assuming a translog term would cut their estimates by a factor of two.

Consider the generic drug (cephalexin) example discussed in Griliches and Cockburn. Roughly speaking the story is as follows: generics enter in at a 50 percent discount and after one year get 60 percent of the quantity market and 43 percent of the revenue share. In this world the incumbent does not change his price and the average price of all versions falls by 0.3. Using Feenstra and Shiells's formula, however, implies not only that the price of generics has declined by 0.5 but that their quality also "improved" (relative to the incumbent brand) by 29 percent! (See table 6C.1.)

Actually the elasticities estimates by Hausman for his functional form and the CES are about the same, but the estimates of consumer's surplus differ widely. My preferred assumption is outside this framework. It allows for heterogeneity of consumers, assumes that the taste for brandedness is distributed uniformly, and implies an average reservation price of $(p_b + p_g)/2$. The resulting total price index, labeled $P(u)$, falls by "only" 22 percent, versus the 40 percent that would be implied by Feenstra and Shiells's formula.

There are a number of problems with Feenstra and Shiells's application: The share of nonpriced items is growing. How nonsubstitutable are they? The implied estimate of $\sigma = 1.9$ is not credible for the average nonpriced item—seven-grain bread versus whole wheat or shirts from Mauritius versus shirts from Singapore.

It is also not well estimated, and the results are very sensitive to that. The approximate standard error for the estimated σ of 2.2 is 1.0. In the formula for the implicit price decline

$$\exp - ds/(\sigma - 1),$$

a difference of one standard deviation would shift the estimate for a 0.10 decline in the share of priced items, from the estimated -0.08, to either -0.50 or -0.05, a rise of 500 percent or a fall of 37 percent in the absolute value of the estimated price change.

There are at least two specification problems associated with the estimated equations: (1) The same excluded-prices story would also apply to domestic goods. I am not sure that the new-goods problem is worse for imports. Internationally traded goods may be *more* standardized than domestic goods. Also,

Table 6C.1 **Branded versus Generics Example**

Period	P_b	P_g	Q_b	Q_g	VS_b	P_{BLS}	P_a	$P(u)$	Feenstra ($\sigma = 2.1$)
1	1		1	0	1	1	1	1	1
2	1	0.5	0.4	0.6	0.57	1	0.7	0.78	0.6

Note: VS_b is the value share of the branded good.

(2) if price is mismeasured, so is the dependent variable, but then their formula for the coefficient becomes $(\beta + 1)(\sigma - 1)$, and the implied $\sigma = 1.2$ is even less credible.

"Aging of lines": Once popular restaurants lose customers over time. We could bring in new ones and make an adjustment for their superiority. But then, some time later, the chefs are hired away and the old restaurants regain their share. Will we come back to the same level? How?

A major finding is that if one allows for the changing mix of import goods this leads to lower estimates of their income elasticity. That makes sense, but how low "should" the import income elasticity be? Can one really explain rising world trade just by the reduction in transport costs and the rising quality of traded goods? I find the notion that traded goods have higher income elasticities quite plausible. The explicit "bias" adjustment to the price index that follows is, however, more problematic. But the advice to collect more data is surely right!

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