Globalization and the Gains from Variety
Author(s): Christian Broda and David E. Weinstein
Source: *The Quarterly Journal of Economics*, Vol. 121, No. 2 (May, 2006), pp. 541-585
Published by: Oxford University Press
Stable URL: http://www.jstor.org/stable/25098800

GLOBALIZATION AND THE GAINS FROM VARIETY*

CHRISTIAN BRODA AND DAVID E. WEINSTEIN

Since the seminal work of Krugman, product variety has played a central role in models of trade and growth. In spite of the general use of love-of-variety models, there has been no systematic study of how the import of new varieties has contributed to national welfare gains in the United States. In this paper we show that the unmeasured growth in product variety from U. S. imports has been an important source of gains from trade over the last three decades (1972–2001). Using extremely disaggregated data, we show that the number of imported product varieties has increased by a factor of three. We also estimate the elasticities of substitution for each available category at the same level of aggregation, and describe their behavior across time and SITC industries. Using these estimates, we develop an exact aggregate price index and find that the upward bias in the conventional import price index over this time period was 28 percent or 1.2 percentage points per year. We estimate the value to U. S. consumers of the expanded import varieties between 1972 and 2001 to be 2.6 percent of GDP.

I. INTRODUCTION

It is striking that in the quarter-century since Krugman [1979] revolutionized international trade theory by modeling how countries could gain from trade through the import of new varieties, no one has structurally estimated the impact of increased variety on aggregate welfare. As a result, our understanding of the importance of new trade theory for national welfare rests on conjecture, calibration, and case studies. While Feenstra [1992], Klenow and Rodriguez-Clare [1997], Bils and Klenow [2001], and Yi [2003] made important inroads into our understanding of the role played by new varieties and differentiated trade, this paper represents the first attempt to answer the question of how much increases in traded varieties matter for the United States. Analyzing the most disaggregated U. S. import data available for the period between 1972 and 2001, we find that consumers have low elasticities of substitution across similar goods produced in dif-

* We would like to thank Fernando Alvarez, Alan Deardorff, Robert Feenstra, Jonathan Eaton, Amartya Lahiri, Mary Amiti, and Kei-Mu Yi for excellent comments and suggestions. Rachel Polimeni provided us with outstanding research assistance. We are grateful to the National Science Foundation for its support of this research (NSF grant No. 0214378). David Weinstein was at the Federal Reserve Bank of New York when most of this research was done. In addition, he would like to thank the Center for Japanese Economy and Business for research support. The views expressed here are those of the authors, and do not necessarily reflect the position of the Federal Reserve Bank of New York, the Federal Reserve System, or any other institution with which the authors are affiliated.

The Quarterly Journal of Economics, May 2006

ferent countries. Moreover, we show that the threefold increase in available varieties arising in the last 30 years has produced a large welfare gain for the United States. We find that consumers are willing to pay 2.6 percent of their income to access the wider set of varieties available in 2001 rather than the set of varieties in 1972. In short, our results provide confirmation of the importance of thinking about international trade within a framework of differentiated goods.

The starting point for our analysis is the seminal work of Feenstra [1994]. In this paper Feenstra develops a robust and easily implementable methodology for measuring the impact of new varieties on an exact price index of a single imported good using only the data available in a typical trade database. Unfortunately, his approach has two drawbacks that have prevented researchers from adopting it more widely. First, it cannot be used to assess the value of the introduction of completely new product categories. Second, Feenstra's methodology tends to generate a large number of elasticities that take on imaginary values, which are hard to interpret. This paper solves both problems and demonstrates the relative ease with which the Feenstra subindexes can be used to compute an *aggregate* price index.

To calculate an aggregate import price index, we first have to estimate a number of parameters. This constitutes our second contribution. In particular, we estimate elasticities of substitution among goods at various levels of aggregation. At the lowest level of aggregation available for trade data (Tariff System of the U.S.A. (TSUSA) seven-digit for 1972–1988 and Harmonized Tariff System (HTS) ten-digit for 1990–2001) we estimate almost 30,000 elasticities. This enables us to directly test a number of stylized facts. For example, we directly demonstrate the validity of Rauch's [1999] conjecture that goods traded on organized exchanges are more substitutable than those that are not. We are able to document that varieties appear to be closer substitutes in more disaggregate product categories. We also find that the median elasticity of substitution has fallen over time indicating that traded goods have become more differentiated.

We then use these estimated parameters to construct a U. S. import price index hewing very closely to theory. Starting with the constant elasticity of substitution utility function which underlies the Spence-Dixit-Stiglitz (henceforth SDS) framework, we compute an exact aggregate price index that allows for changes in varieties. Since this is the same assumption that is used in much of the new trade theory, economic geography, and growth literatures [Helpman and Krugman 1985; Grossman and Helpman 1991; Fujita, Krugman, and Venables 1999], our estimates can be directly applied to these models. Our results suggest that the impact of increased choice on the exact import price index is both statistically and economically significant. Between 1972 and 2001, if one adjusts for new varieties, import prices have been falling 1.2 percentage points per year faster than one would surmise from a conventional price index. If we aggregate across all the years, this means that the variety-adjusted import price index has fallen by 28.0 percent relative to a conventionally measured import price index.

Finally, we are able to use this price decline to obtain an estimate of the gains from new imported varieties under the same structural assumptions as Krugman [1980]. We calculate the compensating variation required for consumers to be indifferent between the set of varieties available in 2001 and that in 1972. We find that consumers are willing to pay 2.6 percent of their income to access the expanded set of varieties available in 2001 rather than the set in 1972. On a per-year basis, this suggests that consumers would pay up to 0.1 percent of their income each year to have access to the net new varieties created that year. We show that the stronger assumptions that are commonly used in the macro literature (e.g., Feenstra [1992], Romer [1994], and Klenow and Rodriguez-Clare [1997]) would lead to welfare gains from variety up to three times larger. Moreover, our results are qualitatively unaffected when we use a Fisher ideal price index or make adjustments for changes in domestic varieties. In sum, our results show that, when measured correctly, increases in imported varieties have had a large positive impact on U. S. welfare.

II. Prior Work

What is a variety? Previous work has not answered this question with a unified voice. In terms of theory, a variety is commonly defined as a brand produced by a firm, the total output of a firm, the output of a country, or the output within an industry in a country.[1] As a result of the variety of definitions of variety, empirical papers are often not strictly comparable. The choices

1. Representative papers would include Hausman [1981], Feenstra [1994], and Roberts and Tybout [1997].

are often driven by data availability and the types of theories that the researchers are examining. While we will make precise our definition of variety later, we want to emphasize that as we discuss prior work, the definition of variety will vary across papers.

Several studies have attempted to measure the impact of new varieties on welfare for individual goods and at the aggregate level. Hausman [1981] pioneered an approach to estimating the gains from new varieties (product line) of an individual good using micro data. He develops a closed-form solution to estimating linear and log-linear demands and calculates the new product's "virtual price," the price that sets its demand to zero. Based on this estimate and on the current price, he calculates the welfare change that results from the price drop of the new product. The advantage of this approach is that by taking enormous care to model, for example, the market for Apple Cinnamon Cheerios, one can obtain extremely precise estimates that can take into account rich demand and supply interactions. However, the data requirements to implement this approach for the tens of thousands of goods that comprise an aggregate price index are simply insurmountable. For this reason, it is not surprising that no one has attempted to estimate aggregate gains from new products using this method.

At the aggregate level, all existing studies rely on calibration or simulation exercises to measure the effect of variety growth. These studies typically define a variety as the imports from a given country or the imports from a given country in a particular industry. They also typically do not focus on how varieties affect prices but rather provide some interesting calculations about potential welfare effects. Feenstra [1992] and Romer [1994], for example, provide numerical exercises showing that the gains from new varieties from small tariff changes can be substantial. Klenow and Rodriguez-Clare [1997] calibrate a model of the impact of trade liberalization on Costa Rica and find only modest gains. They suggest that the low elasticity of substitution and large import shares used in Romer [1994] account for the difference in welfare gains.[2]

These papers have provided an invaluable first step in un-

2. Rutherford and Tarr [2002] simulate a growth model with intermediate input varieties that magnifies the effect of trade liberalization on welfare, and suggest that a 10 percent tariff cut can lead, in the long run, to welfare gains of roughly 10 percent.

derstanding how to move from theory to data, but they require a large number of restrictive simplifying assumptions in order to obtain the estimates. For example, this prior work uses one or at most two elasticities of substitution in order to value varieties. This creates three types of problems. The first arises from assuming that all elasticities of substitution are the same for varieties of different goods. Since presumably consumers care more about varieties of computers than crude oil, it is not clear that all increases in imports correspond to the same gains from increased variety. The second problem arises from assuming that the elasticity of substitution across goods equals that across varieties of a given good. Presumably we care more about the different varieties of fruits than about varieties of apples. The final and perhaps largest problem arises from assuming that all varieties enter into the utility function with a common elasticity. When one is estimating a parameter that is averaging together, say, the impact of an increase of Saudi Arabian oil prices on Mexican oil imports and Japanese car imports, it is hard to interpret the meaning of the elasticity or have intuition for its magnitude.

A different class of problem with calibration exercises stems from the choice of the parameter values and the use of symmetric utility functions (e.g., Romer [1994] and Broda and Weinstein [2004]). Parameter values, such as elasticities of substitution, are often chosen arbitrarily or are estimated from one data set and applied to another data set. An important feature of our study is that all parameters are estimated directly from the relevant data and not chosen in order to obtain sensible values for some other stage of the analysis. Moreover, in the case of a symmetric utility function, since all varieties are valued alike, a count of the number of imported varieties is sufficient to perform welfare calculations. This approach is only valid under the extreme symmetry assumptions underlying the particular utility function used. Indeed, this paper shows that the use of count data, rather than the changes in import volumes as suggested by Feenstra [1994], can be highly misleading as a measure of variety growth if one allows for a more general utility function.

The third problem is related to the way in which previous studies have estimated the single elasticity of substitution. By far the simplest of these approaches is to follow the pioneering work of Anderson [1979] and estimate the elasticity of substitution by regressing bilateral trade flows on various control variables and a measure of trade costs. The coefficient on trade costs is used as the elasticity of substitution among varieties. The major problem with this approach is that one needs to make extreme identifying assumptions in order to ignore simultaneity problems. Chief among these is the assumption that trade costs are completely passed through to consumers. This assumption is almost surely inappropriate for the United States and the other large importers who together account for the majority of world trade. A second problematic identifying assumption is that movements in trade costs are unaffected by movements in import demand. Unfortunately, this assumption will be violated whenever per unit transport costs are a function of import volumes, countries care about import responses when cutting bilateral tariffs, or movements in nontariff barriers are correlated with movements in tariffs. Since all of these conditions are likely to be violated in reality, the estimated elasticities are problematic. Ignoring the simultaneity problem would result in lower estimates of the elasticities of substitution.

Our paper proceeds as follows. In Section III we provide an overview of the basic theoretical contributions on the literature of variety growth and the reasons behind the structure we use in this paper. In Section IV we provide descriptive statistics on the growth in varieties in U. S. imports since 1972. Section V is devoted to the methodology used to compute an exact aggregate price index and to estimate elasticities of substitution that correct for endogeneity bias, measurement error, and that allow for changes in taste and quality parameters. Section VI presents the main results of the paper. We present our conclusions in Section VII.

III. Theory: Why Do Varieties Matter?

All studies that seek to quantify the potential gains from variety are forced to impose some structure on how varieties might affect welfare. Theorists have proposed many ways of modeling this (see, for example, Hotelling [1929], Lancaster [1975], Spence [1976], and Dixit and Stiglitz [1977]), and the assumptions underlying these models are not innocuous. As empirical researchers, we are forced to choose from a number of plausible theories. Our choice of the SDS framework is based on three criteria: prominence, tractability, and empirical feasibility.

There is little question that in international trade, economic geography, and macroeconomics, the SDS framework is the preferred way of specifying how consumers value variety. A major reason for this stems from the tractability of the constant elasticity of substitution (CES) utility function and its close cousin, the Cobb-Douglas. In addition to the work of Krugman, the Dornbusch, Fisher, and Samuelson models are more recent work by Eaton and Kortum [2002] all use CES or Cobb-Douglas functions. Hence, it is quite natural to use this preference structure as the basis of our empirical work. At the very least, our work provides a useful benchmark for thinking about the potential gains from imported varieties within this framework.

A second reason to base our work on the SDS framework is theoretical tractability. As Helpman and Krugman [1985, pp. 124–129] note, preference systems based on the Hotelling and Lancaster models do not easily lend themselves to the creation of aggregate price indexes or utility functions when there is more than one market in the economy. Since one of the main objectives of this study is to build an aggregate price index, we need to use a theoretical structure that will let us aggregate price changes across markets.

Finally, the CES satisfies another important characteristic—empirical feasibility. Demand systems based on CES utility functions are relatively easy to estimate. This is of paramount importance since we need to be able to aggregate estimates of the gains from variety in tens of thousands of markets. Moreover, since we know next to nothing about demand and supply conditions in virtually all of the markets we examine, it is simply not feasible to implement a more complex supply and demand structure.[3] Thus, although one would ideally like to control for all of the complexities present in international markets, the data and time limitations required to perform a careful analysis of all of these markets make this impossible in practice.

Given the way we model how consumers value variety, we now need to be precise about what we mean by a variety. Our reliance on the Krugman [1980] structure might suggest that we adopt a definition of variety that is based on firm-level exports. Unfortunately, there are a number of problems with taking this literal approach to the data. First, by treating all imports from a given firm as a single variety, one may understate the gains from

3. One property of the CES is that, by assumption, consumers care about varieties to some extent. In practice, this assumption does not bias our results because an increase in variety will have a trivial impact on prices and welfare if the estimated elasticity of substitution is large.

variety that occur when a firm starts exporting in more than one product line. Second, it is difficult to obtain bilateral firm-level export data for more than a handful of countries. We therefore opt to use the same definition of variety as in Feenstra [1994]—namely, a seven- or ten-digit good produced in a particular country. To give a concrete example, a good constitutes a particular product, e.g., red wine. A variety, however, constitutes the production of a particular good in a particular country, just as in Armington [1969], e.g., French red wine.

Being clear about this distinction highlights an important difference between monopolistic competition models and comparative advantage models that feature a continuum of goods. Both models share the feature that output of tradables is perfectly specialized in equilibrium. However, they differ in terms of how individual varieties are treated. In the comparative-advantage, continuum-of-goods models, consumers are indifferent about where a good is produced as long as the price does not vary. In other words, these models assume that holding the good fixed, the elasticity of substitution among varieties is infinite. This is in sharp contrast to the Krugman model that hypothesizes that all firms produce differentiated products and hence the elasticity of substitution should be finite.

Despite the sharp theoretical difference, our ability to do precise hypothesis testing is limited. The point estimate for the elasticity of substitution will always be finite, and thus we cannot formally reject this hypothesis. However, by examining the elasticities of substitution at the seven- or ten-digit level, we can obtain a sense of the degree of substitutability among varieties. If the elasticities of substitution tend to be high, say above 10 or 20, then this suggests that the potential for gains from variety, a key theoretical result of the monopolistic competition framework, are small. If they are low, then this suggests that even when we use the most disaggregated trade data in existence, goods are highly differentiated by country. Of course, we cannot rule out that if we had even more disaggregated data, we might find a higher estimate of the elasticity of substitution. Yet even so, we do learn something about the world—at the seven- or ten-digit level of aggregation, it is reasonable to think of goods from different countries as far from perfect substitutes. More importantly for our purposes, low elasticities of substitution across varieties are a necessary condition for increases in the number of varieties to be a source of potential gain.

Turning to welfare, the monopolistic competition model described in Krugman [1979, 1980] suggests two clear channels for the gains from trade arising from variety growth. The first is through reductions in trade costs. If trade costs fall, countries will gain through the import of new varieties.[4] The second is through growth of the foreign country. As the size of the foreign country rises (which in the Krugman framework is equivalent to a rise in its labor force), it will produce more varieties, and this will also be a source of gain for the home country. These gains are in sharp contrast to the gains postulated by comparative advantage models. In these models, all goods are consumed in equilibrium regardless of the level of trade costs or the size of the foreign country. Hence, in comparative advantage models, all gains from reductions in trade costs or increases in the size of a foreign country are achieved through conventional movements in prices and not through changes in the number of goods. One of the distinguishing features of the Krugman model is that a country may gain from trade even though there are no price changes of existing goods.

In sum, although theorists have developed a number of models of variety, our choice of the Dixit-Stiglitz structure stems from that model's prominence, tractability, and empirical implementability. Moreover, since this model can easily explain key stylized facts of how the growth of foreign countries and the reduction of international barriers have contributed to an increase in U. S. imports of varieties, we believe it is a particularly appropriate structure to use in order to obtain estimates of the gains from variety.

IV. DATA: THE GROWTH OF VARIETIES

It is well-known that trade has been growing faster than GDP for many decades. This process, which is a part of what some term "globalization," has had a profound impact on the dependence of the U. S. economy on foreign goods. Over the last 30 years, the share of imports of goods in U. S. GDP has more than

4. The basic Krugman model predicts that a change in tariffs within nonprohibitive values will not change the number of available varieties, although consumers will gain from the falling prices of imported varieties. Romer [1994], however, presents a simple extension of this model to allow for fixed costs of accessing foreign markets so that the number of available varieties rises with a fall in tariffs.

doubled: rising from 4.8 percent in 1972 to 11.7 percent in 2001.[5] The causes for this explosion in trade stem from a number of sources that have been explored in a vast literature. Most studies attribute the source of the rise to three interrelated causes: reductions in trade costs, relaxations of capital controls (e.g., barriers to foreign direct investment), and the relative growth of many East Asian and other economies outside of the United States.

This rise in U. S. imports has been accompanied by a rise in another phenomenon that has received much less attention—a dramatic rise in imported varieties. Table I gives a preliminary overview of the extent of this increase. Between 1972 and 1988 we rely on the TSUSA seven-digit data and in later years on the ten-digit HTS data [Feenstra 1996; Feenstra, Hanson, and Lin 2004]. We define a good to be a seven- or ten-digit category, and, as mentioned in the previous section, a variety is defined as the import of a particular good from a particular country.[6] We do not report numbers for 1989 because the unification of Germany means that data for that year are not comparable with later HTS data.[7]

Using our definition of varieties, Table I reports that in 1972 the United States imported 71,420 varieties (i.e., 7731 goods from an average of 9.2 countries) and in 2001 there were 259,215 varieties (16,390 goods from an average of 15.8 countries). Ultimately, we will want to make comparisons across years, and to do that properly we will need to formally deal with a host of issues relating to whether the data for two different years are truly comparable. For now, we put these issues aside and focus on the crude measure of variety that we can glean from the sample statistics.

The second column of Table I reports the number of goods for which imports exceeded one dollar in a given year. There are two features of this column that are important to note here. First, comparing the values for 1988 and 1990, there appears to be little

5. Data are from the World Bank *World Development Indicators* unless otherwise stated.

6. This definition matters less than one might suppose for our later empirical work since we will estimate elasticities of substitution across varieties of a good and let the data tell us how important differences among varieties are. For the time being, however, we will leave aside the question of how substitutable goods produced in different countries are, and simply focus on the number of varieties.

7. All the countries in the former Soviet Union are aggregated together throughout our analysis.

TABLE I
VARIETY IN U. S. IMPORTS (1972–2001)

U. S. Imports 1972–1988

	Year	Number of TSUSA categories	Median number of exporting countries	Average number of exporting countries	Total number of varieties (country-good pairs)	Share of total U. S. imports in year
	(1)	(2)	(3)	(4)	(5)	(6)
All 1972 goods	1972	7731	6	9.2	71420	1.00
All 1988 goods	1988	12822	9	12.2	156669	1.00
Common 1972–1988	1972	4167	6	8.4	35060	0.41
Common 1972–1988	1988	4167	10	12.2	50969	0.33
1972 not in 1988	1972	3553	7	10.2	36355	0.59
1988 not in 1972	1988	8640	8	12.7	105696	0.67

U. S. Imports 1990–2001

	Year	Number of HTS categories	Median number of exporting countries	Average number of exporting countries	Total number of varieties (country-good pairs)	Share of total U. S. imports in year
	(1)	(2)	(3)	(4)	(5)	(6)
All 1990 goods	1990	14572	10	12.5	182375	1.00
All 2001 goods	2001	16390	12	15.8	259215	1.00
Common 1990–2001	1990	10636	10	12.4	132417	0.73
Common 1990–2001	2001	10636	13	16.3	173776	0.67
1990 not in 2001	1990	3936	10	12.7	49958	0.27
2001 not in 1990	2001	5754	11	14.8	85439	0.33

Source: NBER CD-ROM and http://data.econ.ucdavis.edu/international/usixd/wp5514d.html

difference in the number of categories with positive imports in the TSUSA and HTS systems. Second there appears to be a dramatic increase in the number of U. S. import categories over the time periods. Combining the increases over the periods 1972–1988 and 1990–2001, it appears that the number of good categories almost doubled. This establishes the importance of thinking about real or apparent new goods or categories when calculating changes in import structure and the price of imports.

Columns (3) and (4) report the median and average number of countries exporting a good to the United States. These data also reveal a substantial increase in the number of countries supplying each individual good. Between 1972 and 2001 the median number of countries doubled, rising from six countries in 1972 to twelve countries today. Similarly, the average number of countries rose 30 percent between 1972 and 1988 and another 30 percent in between 1990 and 2001, resulting in an aggregate increase of 67 percent. In other words, even if we leave aside the issue of why the number of imported categories has increased over time, the data reveal that there has been a dramatic increase in the number of countries supplying each individual good.

This effect can also be seen if we restrict ourselves to the set of goods that were imported at the start and end of each sample period. In rows 3 and 4 of Table I, we present data on the set of common goods within each sample. The data reveal that the increase in countries supplying these goods was, if anything, even more pronounced than the increase for the sample as a whole. The aggregate increase in the median number of countries supplying common goods was 117 percent, and the average rose 91 percent.

The last two lines provide sample statistics for the set of categories that ceased to exist or appeared during this time period. Roughly a third to a half of the categories in which the United States recorded positive imports at the start of either period did not contain positive imports at the end of the period. Similarly, somewhere between a third and two-thirds of the products imported at the end of each period were not imported at the start of the sample. Once again, we will have to return to the question of whether this represents the actual birth and death of products or simply product categories, but the table underscores that there are substantial changes in the measured composition of imports across both time periods.

Taken together, the data in Table I suggest that the number of varieties rose 119 percent in the first period and 42 percent in the second period—a total increase of 212 percent. This increase constitutes more than a threefold increase in the number of varieties over the last three decades. Roughly half of this increase appears to have been driven by an increase in the number of goods and half by an increase in the number of countries supplying each good.

The fact that the number of countries supplying each good almost doubled serves as prima facie evidence of a startling increase in the number of varieties. The most plausible explanations for this rise involve some story of the globalization process coupled with an assumption that goods are differentiated by country (as in Krugman [1980], Romer [1994], and Rutherford and Tarr [2002]). For example, reductions of trade costs may have made it cheaper to source new varieties from different countries. Alternatively, the growth of economies like China, Korea, and India has meant that they now produce more varieties that the United States would like to import. But, of course, if these goods are differentiated by country, then this implies that there must be some gain from the increase in variety—a point that we will address in the next section.

One can obtain a better sense of the forces that have been driving the increase in variety if we break the data up by exporting country. Table II presents data on the numbers of goods exported to the United States by country. The first column ranks them from highest to lowest for 1972, and the following columns rank them for subsequent years. Not surprisingly, the countries that export the most varieties to the United States tend to be large, high-income, proximate economies. Looking at what has happened to the relative rankings over time, however, reveals a number of interesting stylized facts. First, Canada and Mexico have risen sharply in the rankings. Canada moved from being the fourth largest source of varieties to first place, while Mexico moved from thirteenth to eighth place. This may reflect free trade areas and other trade liberalizations between the United States and these countries over the last several decades.

Growth, perhaps coupled with liberalization, also appears to have played some role. Fast growing economies like China and Korea rose dramatically in the rankings. For example, in 1972 China only exported 710 different goods to the United States as opposed to 10,315 in 2001. This fourteen-fold increase in the number of varieties produced a dramatic change in China's relative position: moving from the twenty-eighth most important source of varieties in 1972 to the fourth most important today. Similarly, after India began its period of liberalization in the last decade, its growth rate rose sharply as did the number of goods it began exporting. At the other extreme, economies like Japan and

TABLE II
RANKING IN TERMS OF NUMBER OF GOODS IMPORTED BY THE UNITED STATES

Country	Ranking in year: 1972	1988	1990	2001
Japan	1	1	3	7
United Kingdom	2	4	4	3
Germany	3	3	2	2
Canada	4	2	1	1
France	5	6	5	6
Italy	6	5	6	5
Switzerland	7	11	11	11
Hong Kong	8	9	12	16
Netherlands	9	13	13	14
Taiwan	10	7	7	9
Spain	11	14	15	12
Belgium-Luxemburg	12	15	14	15
Mexico	13	12	10	8
Sweden	14	17	16	19
Denmark	15	22	21	23
Austria	16	18	18	21
India	17	19	23	13
Rep. of Korea	18	8	9	10
Brazil	19	16	17	18
Australia	20	20	20	20
Israel	21	21	22	22
Portugal	22	26	28	32
Norway	23	31	31	37
Ireland	24	27	26	28
Finland	25	28	30	31
Colombia	26	33	34	35
Philippines	27	25	25	26
China	28	10	8	4
Argentina	29	29	29	39
Greece	30	38	44	47

Top 30 countries in 1972 included. Same notes as in Table I apply.

Argentina have seen fairly substantial drops in the relative number of varieties they export.

The importance of these countries for the growth in available U. S. varieties can be seen in Table III. The first column presents the ratio of the net change in varieties between 1972 and 1988 from a given country to the change in varieties entering the United States as a whole. The second column reports the average share of imports from that country in the first time periods. The

TABLE III
COUNTRY CONTRIBUTION TO GROWTH IN U. S. VARIETIES (1972–1988/1990–2001)

Country	Contribution 1972–1988	Average share of U. S. imports (*)	Country	Contribution 1990–2001	Average share of U. S. imports (*)
China	4.8%	1.0%	China	5.7%	6.0%
Taiwan	4.4%	4.0%	India	4.4%	0.7%
Rep. of Korea	4.4%	2.9%	Mexico	3.7%	8.8%
Canada	4.2%	22.7%	Spain	2.9%	0.6%
Italy	4.0%	2.9%	South Africa	2.6%	0.4%
Germany	3.8%	6.9%	Italy	2.6%	2.3%
France	3.8%	2.6%	Indonesia	2.5%	0.8%
Japan	3.6%	18.4%	Canada	2.5%	18.7%
United Kingdom	3.5%	4.7%	Turkey	2.3%	0.3%
Hong Kong	3.1%	2.3%	Thailand	2.3%	1.2%
Mexico	3.0%	4.0%	Australia	2.1%	0.7%
Switzerland	2.6%	1.1%	France	2.1%	2.7%
Brazil	2.6%	1.9%	Rep. of Korea	2.0%	3.4%
Netherlands	2.2%	1.1%	Belgium-Luxemburg	1.9%	0.9%
Thailand	2.2%	0.5%	Poland	1.8%	0.1%
Singapore	1.9%	1.1%	Malaysia	1.8%	1.5%

A U. S. variety is defined as a TSUSA-exporting country pair in 1972–1988 and HTS-exporting country pair in 1990–2001. (*) Log ideal weights used as average shares (see text for a definition).

third and fourth columns repeat this exercise for the second time period. The table highlights the importance that industrializing Asia has played in the creation of new varieties. Particularly prominent is the role played by China. In the first period, China accounted for almost 5 percent of aggregate U. S. variety growth, even though China only accounted for an average of 1 percent of U. S. imports. Other rapidly growing or liberalizing countries, such as Taiwan, Korea, India, and Mexico, also contributed heavily to the increase in available varieties.

One simple way of numerically exploring the association between growth, trade, and the number of varieties exported to the United States is to use a strategy similar to that in Eaton, Kortum, and Kramarz [2004]. Let M_{et} be the total imports from exporting country e to the United States in time t, and V_{et} the number of varieties of goods imported by exporting country e. We can use the following identity, $M_{et} \equiv V_{et}\bar{m}_{et}$, where $\bar{m}_{et}$ is the average exports per variety, to describe the relationship between the two variables. Moreover, it is standard to model bilateral trade using the gravity model, where bilateral imports are a function of the market sizes of the United States and the exporting country, and measures of geographic barriers between the two countries. That is, $M_{et} = \zeta Q_e Q_{us}/d_{us,e}$, where Q_e is the output of country e, ζ is a constant, and $d_{us,e}$ a measure of the distance between countries.

From these two relationships we can derive and estimate the following regression,

$$\ln V_{e,01} - \ln V_{e,72} = \alpha + \beta(\ln Q_{e01} - \ln Q_{e72}) + \gamma(\ln \vartheta_{e01} - \ln \vartheta_{e72}) + \varepsilon_e,$$

where $\vartheta_{et} = \zeta Q_{us}/d_{us,e} = M_{et}/Q_e$ measures exports of country e to the United States as a share of its GDP and α, β, and γ are parameters to be estimated. This is the time-series analog of the regression in Eaton, Kortum, and Kramarz [2004]. This regression yields the following coefficients: $\beta = 0.32$ (0.04), $\gamma = 0.35$ (0.04), where robust standard errors appear in brackets, and $R^2 = 0.55$.[8] These results suggest that holding fixed a country's share of exports to the United States, a 1 percent increase in trading partner's size is associated with a 32 percent increase in the number of varieties exported to the United States. Similarly, a 1 percent increase in imports by the United States from a country is associated with 35 percent more varieties exported from that country. These results further suggest that the increase in varieties was not random. Rather, foreign countries that exported and grew more tended to disproportionately increase the number of varieties they exported to the United States.

We will now formally deal with how to correctly measure and value these increases in product varieties. In particular, we discuss how the methodology used is robust to a host of issues that have been ignored in this descriptive section.

V. Empirical Strategy

V.A. *The Feenstra Price Index*

In this subsection we extend Feenstra's [1994] derivation of the exact price index of a single CES aggregate good that allows for both new varieties and taste or quality changes in existing

8. When a similar regression is run on the cross section for the year 2001, we obtain the following parameters, $\beta = 0.69$ (0.03), $\gamma = 0.42$ (0.05), and $R^2 = 0.83$. Eaton, Kortum, and Kramarz [2004] using export firm data from France to the rest of the world, get the following coefficients: $\beta = 0.62$ (0.02), $\gamma = 0.88$ (0.03), and $R^2 = 0.90$.

varieties, to the case of several CES aggregate goods. Our objective in this subsection is to build an exact aggregate price index corresponding to a CES utility function that can be used to evaluate the impact of variety changes on prices and welfare. The first step toward deriving an aggregate exact price index is to define the utility function. Suppose that the preferences of a representative agent can be denoted by a three-level utility function (similar to Helpman and Krugman [1985, Ch. 6]) that aggregates imported varieties into composite imported goods, then aggregates these imported goods into a composite import good, and finally combines this imported good with the domestic good to produce utility. For expositional purposes, we begin by specifying the upper level utility function as

$$U_t = (D_t^{(\kappa-1)/\kappa} + M_t^{(\kappa-1)/\kappa})^{\kappa/(\kappa-1)}; \quad \kappa > 1, \tag{1}$$

where M_t is the composite imported good to be defined below, D_t is the domestic good, and κ is the elasticity of substitution between both goods. While this functional form allows the share of imports to vary over time, it creates a certain degree of separability between imports and domestic goods that will prove useful as we develop our price index. This assumption will be important for calculating an aggregate import price index.

Moving to the second tier, we define the composite imported good as

$$M_t = \left(\sum_{g \in G} M_{gt}^{(\gamma-1)/\gamma} \right)^{\gamma/(\gamma-1)}; \quad \gamma > 1, \tag{2}$$

where M_{gt} is the subutility derived from the consumption of imported good g in time t, γ denotes the elasticity of substitution among imported goods, and G is the set of all imported goods.

A particularly useful form of M_{gt} is the *nonsymmetric* CES function, which can be represented by

$$M_{gt} = \left(\sum_{c \in C} d_{gct}^{1/\sigma_g} (m_{gct})^{(\sigma_g-1)/\sigma_g} \right)^{\sigma_g/(\sigma_g-1)}; \quad \sigma_g > 1 \ \forall g \in G, \tag{3}$$

where σ_g is the elasticity of substitution among varieties of good g, which is assumed to exceed unity; for each good, imports are treated as differentiated across countries of supply, c (as in Armington [1969]).[9] That is, we identify varieties of import good g with their countries of origin. C is the set of *all* countries; d_{gct} denotes a taste or quality parameter for good g from country c.

Let $I_{gt} \subset C$ be the subset of all varieties of good g consumed in period t. The minimum unit-cost function of subutility function in (3) is given by the following expression:

$$\phi_{gt}^{M}(I_{gt},\mathbf{d}_{gt}) = \left(\sum_{c \in I_{gt}} d_{gct}(p_{gct})^{1-\sigma_g} \right)^{1/(1-\sigma_g)}, \tag{4}$$

where p_{gct} is the price of variety c of good g in period t and $\mathbf{d}_{gt}$ is the vector of taste or quality parameters for each country. Note that (4) can be used to illustrate the essence of the love-of-variety approach and the source of deficiencies in conventional price indices. Suppose that V_g varieties of good g are available to consumers and that $d_{gc} = 1\ \forall c \in C$ (i.e., M_g is symmetric). Then in a standard monopolistic competition model all varieties will be equally priced at p_g. In this case, the minimum unit-cost function becomes $\phi_g^M = V_g^{1/(1-\sigma_g)}p_g$. For a given p_g, an increase in V_g implies that the minimum cost required to attain a given level of utility falls. However, a conventional price index that is based only on common varieties will not capture the fall in minimum unit-costs, or equivalently, the rise in utility.

The minimum unit-cost function of (2), in turn, can be denoted by

$$\phi_t^M = \left(\sum_{g \in G} (\phi_{gt}^{M}(I_{gt},\mathbf{d}_{gt}))^{1-\gamma} \right)^{1/(1-\gamma)}. \tag{5}$$

And the overall price index is given by

$$p_t = [(p_t^D)^{1-\kappa} + (\phi_t^M)^{1-\kappa}]^{1/(1-\kappa)}, \tag{6}$$

where the price of the domestic good is given by p_t^D. Equations (4)–(6) constitute the main building blocks for the calculation of exact aggregate price indices.

We turn next to the derivation of the aggregate bias generated by ignoring new varieties. We proceed in three steps: first,

9. One of the features and limitations of the CES functional form is that the elasticity of substitution plays a dual role as a measure of substitution across varieties and a key factor in evaluating new varieties. This functional form assumption makes the CES attractive for theoretical and empirical researchers, but one can contemplate more complex relationships. Brown, Deardorff, and Stern [1996] calibrate a model with variety growth using a more general CES function.

we review Feenstra's [1994] contribution, namely, to generalize the exact price index of a single good to the case of new and disappearing product varieties; second, we derive the *aggregate* exact import price index for (2); and last, we provide a description of the useful properties of this aggregate index.

Diewert [1976] defines an exact price index for good g over a constant set of varieties as

$$(7)\qquad P_g^M(\mathbf{p}_{gt},\mathbf{p}_{gt-1},\mathbf{x}_{gt},\mathbf{x}_{gt-1},I_g) = \frac{\phi_{gt}^M(I_g,\mathbf{d}_g)}{\phi_{gt-1}^M(I_g,\mathbf{d}_g)},$$

where $I_g = I_{gt} \cap I_{gt-1}$ is the set of varieties consumed in periods t and $t-1$, and taste parameters are constant over time, $d_{gct} = d_{gct-1} = d_{gc}$ for $c \in I_g$. Since in this case varieties are constant over time, $I_g = I_{gt} = I_{gt-1}$. $\mathbf{x}_{gt}$ and $\mathbf{x}_{gt-1}$ are the cost-minimizing quantity vectors of good g's varieties given the prices of all varieties, $\mathbf{p}_{gt}$ and $\mathbf{p}_{gt-1}$. This means that an exact price index has the salient feature that a change in the index exactly matches the change in minimum unit-costs.[10] As noted by Diewert, a remarkable feature of (7) is that the price index does not depend on the unknown quality parameters d_{gc}. The intuition for this result is that all of the information contained in the quality parameters is captured by the levels of consumption.

In the case of the CES unit-cost function, Sato [1976] and Vartia [1976] have derived its exact price index to be

$$(8)\qquad P_g(\mathbf{p}_{gt},\mathbf{p}_{gt-1},\mathbf{x}_{gt},\mathbf{x}_{gt-1},I_g) = \prod_{c\in I_g}\left(\frac{p_{gct}}{p_{gct-1}}\right)^{w_{gct}}.$$

This is the geometric mean of the individual variety price changes, where the weights are ideal log-change weights.[11] These weights are computed using cost shares s_{gc} in the two periods, as follows:

$$(9)\qquad s_{gct} = \frac{p_{gct}x_{gct}}{\sum_{c\in I_g} p_{gct}x_{gct}}$$

10. Diewert [1976] also presents the dual of (7), where the exact quantity index has to match the change in utility from one period to the other.

11. As explained in Sato [1976], a price index P that is dual to a quantum index, Q, in the sense that $PQ = E$ and shares an identical weighting formula with Q is defined as "ideal." Fisher [1922] was the first to use the term ideal to characterize a price index. He noted that the geometric mean of the Paasche and Laspeyres indices is ideal.

$$(10) \qquad w_{gct} = \frac{(s_{gct} - s_{gct-1})/(\ln s_{gct} - \ln s_{gct-1})}{\sum_{c \in I_g} ((s_{gct} - s_{gct-1})/(\ln s_{gct} - \ln s_{gct-1}))}.$$

The numerator of (10) is the difference in cost shares over time divided by the difference in logarithmic cost shares over time.

The exact price index of good g in (7), P_g, requires that all varieties be available in the two periods. Feenstra [1994] showed how to modify this exact price index for the case of different, but overlapping, sets of varieties in the two periods. Proposition 1 states Feenstra's main theoretical contribution, the relationship between the conventional price index and the exact price index that incorporates changes in variety for a single good.

PROPOSITION 1.[12] For $g \in G$, if $d_{gct} = d_{gct-1}$ for $c \in I_g = (I_{gt} \cap I_{gt-1})$, $I_g \neq \varnothing$, then the exact price index for good g with change in varieties is given by

$$(11) \quad \pi_g(\mathbf{p}_{gt},\mathbf{p}_{gt-1},\mathbf{x}_{gt},\mathbf{x}_{gt-1},I_g) = \frac{\phi^M_{gt}(I_{gt},\mathbf{d}_g)}{\phi^M_{gt-1}(I_{gt-1},\mathbf{d}_g)}$$
$$= P^M_g(\mathbf{p}_{gt},\mathbf{p}_{gt-1},\mathbf{x}_{gt},\mathbf{x}_{gt-1},I_g)\left(\frac{\lambda_{gt}}{\lambda_{gt-1}}\right)^{1/(\sigma_g-1)},$$

where

$$\lambda_{gt} = \frac{\sum_{c \in I_g} p_{gct}x_{gct}}{\sum_{c \in I_{gt}} p_{gct}x_{gct}} \quad \text{and} \quad \lambda_{gt-1} = \frac{\sum_{c \in I_g} p_{gct-1}x_{gct-1}}{\sum_{c \in I_{gt-1}} p_{gct-1}x_{gct-1}}.$$

This result states that the exact price index with variety change (i.e., $\pi_g(I_g)$ for short) is equal to the "conventional" price index, $P^M_g(I_g)$ (i.e., the exact price index of the common varieties over time), multiplied by an additional term, $(\lambda_{gt}/\lambda_{gt-1})^{1/(\sigma_g-1)}$, which captures the role of the new and disappearing varieties.[13] Note that λ_{gt} equals the fraction of expenditure in the varieties that are available in both periods (i.e., $c \in I_g = (I_{gt} \cap I_{gt-1})$) relative to the entire set of varieties available in period t (i.e., $c \in I_{gt}$). Thus, this additional term implies that the higher the expenditure share of new varieties, the lower is λ_{gt}, and the smaller is the exact price index relative to the conventional price index. In the

12. The appendix of Feenstra [1994] provides the proof of a more general proposition, where $c \in I_g \subseteq (I_{gt} \cap I_{gt-1})$.
13. All of the index numbers used in this paper suffer from the classic "index number problem." In particular, results are dependent on the base year or years used. Since we are examining long-run changes, we use two base years 1972 and 1990.

symmetric case, (11) simply becomes $\pi_g(I_g) = P_g^M(I_g)(V_{gt-1}/V_{gt})^{1/(\sigma_g-1)}$, where V_{gt-1} and V_{gt} are the number of varieties of good g consumed in periods t and $t-1$, respectively. It is easy to see that, in this case, an increase in the number of varieties leads directly to a fall in the exact price index relative to the conventional price index.

The Feenstra price index also depends on the good-specific elasticity of substitution, σ_g. As σ_g grows, the term $1/(\sigma_g - 1)$ approaches zero, and the bias term $(\lambda_{gt}/\lambda_{gt-1})^{1/(\sigma_g-1)}$ becomes unity. That is, when existing varieties are close substitutes to new or disappearing varieties, changes in variety will not have a large effect on the exact price index. By contrast, when σ_g is close to unity, varieties are not close substitutes, $1/(\sigma_g - 1)$ is high, and therefore new varieties are very valuable, and disappearing varieties are very costly.

Having derived the exact price index with variety change for the subutility function in (3), we can now obtain the *aggregate* exact import price index for (2) which is summarized in the following proposition.[14]

PROPOSITION 2. If $I_g \neq \varnothing\ \forall g \in G$ and $d_{gct} = d_{gct-1}$ for $c \in I_g$ $\forall g \in G$, then the exact *aggregate* import price index with variety change is given by

$$(12) \qquad \Pi^M(\mathbf{p}_t,\mathbf{p}_{t-1},\mathbf{x}_t,\mathbf{x}_{t-1},I) = \frac{\phi_t^M(I_t,\mathbf{d})}{\phi_{t-1}^M(I_{t-1},\mathbf{d})} = CIPI(I) \prod_{g\in G} \left(\frac{\lambda_{gt}}{\lambda_{gt-1}}\right)^{w_{gt}/(\sigma_g-1)},$$

where $CIPI(I) = \Pi_{g\in G} P_g(I_g)^{w_{gt}}$ and w_{gt} are log-change ideal weights.

The second equality follows from applying (8) to the CES bundle of imported goods, and Samuelson [1965]. By replacing (7) and (11) in (12), we obtain the relationship between the *aggregate* conventional import price index (*CIPI*) and the *aggregate* exact import price index. For future reference, we will refer to the geometric weighted average of the λ ratios in (12) as the *aggregate* import bias that results from ignoring new varieties in all product categories. The empirical measurement of this aggregate bias is

14. As will be clear in the empirical section, the number of goods, G, is assumed constant over time.

the focus of the empirical section that follows and represents the main contribution of this paper. The overall price index Π is then given by

$$\Pi = \left(\frac{p_t^D}{p_{t-1}^D}\right)^{w_t^D} (\Pi^M)^{w_t^M}, \tag{13}$$

where the exponents are the log-change ideal weights.

Basing an aggregate import price index on (12) corrects a host of problems that plagued prior work. First, this framework allows varieties to account for different shares of expenditures due to quality or taste differences. This contrasts with prior work measuring variety growth, which replace our lambda ratio, $\lambda_{gt}/\lambda_{gt-1}$, with the ratio of the number of varieties in each period, V_{gt-1}/V_{gt} (e.g., Romer [1994] and Broda and Weinstein [2004]). As equation (11) suggests, replacing the lambda ratio with the ratio of the number of varieties in the two periods can yield substantial biases. These "quality biases" can be quite large. For example, if new varieties represent only a small (large) share of the total expenditure in a good, then a simple count of varieties will grossly overestimate (underestimate) the true impact of new varieties.

Second, this framework eliminates the "symmetry bias" that arises from assuming that all varieties are interchangeable. As equation (12) indicates, the correct price index should allow for elasticities of substitution among varieties of different goods to vary. This implies that the same increase in price of a variety of two different goods may be valued differently by consumers. Thus, measuring the aggregate bias requires that these elasticities of substitution be estimated (this is the focus of the next section). In other words, we do not require that $w_{gt}(G)$ or σ_g be the same for all goods. Moreover, since we have a three-tiered CES utility function, we do not require that the elasticity of substitution among varieties be the same as that across goods.[15]

Third, the aggregate price index in (12) is robust to a wide variety of data problems arising from the creation and destruction of product categories g. For example, if goods are randomly

15. Proposition 2 still holds if the existing *goods* are bundled into CES aggregates with different elasticities between them rather than using a common elasticity γ as in (2).

split or merged, then the index remains unchanged.[16] By contrast, a measure based on the number of varieties would erroneously register a fall or rise in the price level. Similarly, it can be shown that if categories are split when a product category becomes large and merged when it becomes small, then the index also will remain unchanged.[17] Finally our index is also robust to the possibility that there may be more than one variety contained in the imports from a given country of a seven- or ten-digit good.[18]

V.B. *Identification and Estimation of the Elasticity of Substitution*

In order to estimate the impact of new imported varieties on the price index, we first need to obtain estimates of the elasticity of substitution between varieties of each good. In this subsection we present a simple model of import demand and supply equations to estimate this elasticity of substitution. Our estimation procedure closely resembles the approach in Feenstra [1994], except that we supplement it by allowing for a more general estimation technique and extend his treatment of measurement error. We depart from the usual gravity equation model to estimate elasticities of substitution in that we allow for an upward sloping export supply curve. The estimation procedure we use allows for random changes in the taste parameters of imports by country and is robust to measurement error from using unit values that are not proper price indices.

The import demand equation for each variety of good g can be derived from the utility function in (3). Expressed in terms of

16. A simple example can help understand the intuition of this result. Assume that there are two varieties (1 and 2) of good g in period $t - 1$, and $p_{g1t-1}q_{gt-1} = p_{g2t-1}q_{g2t-1} = 5$. In period t the consumption of variety 1 remains unchanged, but variety 2 splits into varieties 3 and 4, and consumption is given by $p_{g2t}q_{g2t} = 0$, $p_{g3t}q_{g3t} = 2$, $p_{g4t}q_{g4t} = 3$. It is easy to show that our measure of the price movement arising from new varieties, $(\lambda_{gt}/\lambda_{gt-1})^{1/(\sigma_g-1)}$, is unaffected (as it should be). Similarly, we can show that if the number of goods categories increases, our index will not change. Note also that if the number of varieties were used instead of the shares, the index would fall from 1 to 2/3.

17. The proof is available from the authors. It is possible to have a bias if statistical agencies split categories that grow but never destroy old categories. However, if this were true, we should observe the average imports per category falling over time rather than the relatively constant size of categories that we actually see.

18. Feenstra [1994] shows that the effects of multi-variety per product-country pair acts in the same way as a change in the taste parameter or quality parameter for that country's imports.

shares and changes over time, the equation for the import demand of a particular variety is the following:[19]

$$\Delta \ln s_{gct} = \varphi_{gt} - (\sigma_g - 1)\Delta \ln p_{gct} + \varepsilon_{gct}, \tag{14}$$

where $\varphi_{gt} = (\sigma_g - 1)\ln[\phi^M_{gt}(d_t)/\phi^M_{gt-1}(d_{t-1})]$ is a random effect as b_t is random and $\varepsilon_{gct} = \Delta \ln d_{gct}$. As opposed to most of the empirical literature that uses a gravity model to estimate the elasticity of substitution which implicitly assumes a horizontal supply curve (and therefore, no simultaneity bias), we allow the export supply equation of variety c to vary with the amount of exports. The export supply equation is given by the following expression:

$$\Delta \ln p_{gct} = \psi_{gt} + \frac{\omega_g}{1 + \omega_g} \Delta \ln s_{gct} + \delta_{gct}, \tag{15}$$

where $\varphi_{gt} = -\omega_g \Delta \ln E_{gt}/(1 + \omega_g)$, $\omega_g \geq 0$ is the inverse supply elasticity (assumed to be the same across countries) and $\delta_{gct} = \Delta \ln \upsilon_{gct}/(1 + \omega_g)$ captures any random changes in a technology factor υ_{gct}. Note that $\omega_g = 0$ is a special case of (15), where the supply curve is horizontal and there is no simultaneity bias. More importantly for the identification strategy is that we assume that $E(\varepsilon_{gct}\delta_{gct}) = 0$. That is, once good-time specific effects are controlled for, demand and supply errors at the variety level are assumed to be uncorrelated.

As in Feenstra, it is convenient to write (14) and (15) in a way that φ_{gt} and ψ_{gt} are eliminated so that we can use the assumption that error terms are independent across equations. For this reason, we choose a reference country k and differences demand and supply equations denoted in equations (14) and (15) relative to country k:

$$\Delta^k \ln s_{gct} = -(\sigma_g - 1)\Delta^k \ln p_{gct} + \varepsilon^k_{gct} \tag{16}$$

$$\Delta^k \ln p_{gct} = \frac{\omega_g}{1 + \omega_g} \Delta^k \ln s_{gct} + \delta^k_{gct}, \tag{17}$$

where $\Delta^k x_{gct} = \Delta x_{gct} - \Delta x_{gkt}$, $\varepsilon^k_{gct} = \varepsilon_{gct} - \varepsilon_{gkt}$ and $\delta^k_{gct} = \delta_{gct} - \delta_{gkt}$. To take advantage of $E(\varepsilon^k_{gct}\delta^k_{gct}) = 0$, we multiply (16) and (17) to obtain

$$(\Delta^k \ln p_{gct})^2 = \theta_1(\Delta^k \ln s_{gct})^2 + \theta_2(\Delta^k \ln p_{gct}\Delta^k \ln s_{gct}) + u_{gct}, \tag{18}$$

19. We use shares (s_{gct}) rather than quantities because shares should not be influenced by the measurement error unit values [Kemp 1962].

where

$$\theta_1 = \frac{\omega_g}{(1+\omega_g)(\sigma_g - 1)}, \quad \theta_2 = \frac{1 - \omega_g(\sigma_g - 2)}{(1+\omega_g)(\sigma_g - 1)}$$

and

$$u_{gct} = \varepsilon^k_{gct}\delta^k_{gct}.$$

Unfortunately, $\beta_g = \binom{\sigma_g}{\omega_g}$ cannot be consistently estimated from (18) as the error term u_{gct} is correlated with the regressands that depend on prices and expenditure shares. However, it is still possible to obtain consistency by exploiting the panel nature of the data set combined with the assumption that demand and supply elasticities are constant over varieties of the same good. In particular, we can define a set of moment conditions for each good g, by using the independence of the unobserved demand and supply disturbances for each country over time; that is,

$$G(\beta_g) = E_t(u_{gct}(\beta_g)) = 0 \quad \forall c. \tag{19}$$

As long as all countries exporting good g satisfy the following condition:

$$\chi^2_{\varepsilon^k_{gc}}/\chi^2_{\varepsilon^k_{gc}} \neq \chi^2_{\delta^k_{gc}}/\chi^2_{\delta^k_{gc}},$$

where χ^2_x is the variance of x, equation (19) implies having V_g independent moment conditions for each good to estimate the two parameters of interest.[20] This condition effectively implies that the regressands between the two countries c and c' are not collinear which would not let us solve the identification problem faced in equation (18). This condition is formally derived in Feenstra [1994].

For each good g, all the moment conditions that enter the GMM objective function can be combined to obtain Hansen's [1982] estimator:

$$\hat{\beta}_g = \arg\min_{\beta \in B} G^*(\beta_g)' W G^*(\beta_g), \tag{20}$$

where $G^*(\beta)$ is the sample analog of $G(\beta)$, W is a positive definite

20. Rigobon [2003] shows that by using the heteroskedasticity of one of the endogenous variables he can achieve full identification. In particular, he identifies the desired coefficients by dividing the sample into periods of high and low volatility and constraining the parameters and variances that are allowed to change across periods. Our approach is analogous in that we require that the relative variances must differ across countries for a given time period.

weighting matrix to be defined below, and B is the set of economically feasible β (i.e., $\sigma_g > 1$; $\omega_g > 0$). We implement this estimator by first estimating the between estimates of θ_1 and θ_2 and then solving for β_g as in Feenstra [1994]. If this produces imaginary estimates or estimates of the wrong sign, we use a grid search of β's over the space defined by B. In particular, we evaluate the GMM objective function for values of $\sigma_g \in [1.05,131.5]$ at intervals that are 5 percent apart.[21] Standard errors for each parameter were obtained by bootstrapping the grid-searched parameters.

The problem of measurement error in unit values motivates our weighting scheme. In particular, there is good reason to believe that unit values calculated based on large volumes are much better measured than those based on small volumes of imports. In the appendix we show that this requires us to add one additional term inversely related to the quantity of imports from the country and weight the data so that the variances are more sensitive to price movements based on large shipments than small ones. The use of the between estimate coupled with our need to estimate σ_g, ω_g, and a constant means that we need data from at least three countries to identify β.

VI. Results

Our estimation strategy involves four stages. First, we need to obtain estimates of the elasticity of substitution, σ_g, by estimating (20). Second, we obtain estimates of how much variety

21. To make sure that we were using a sufficiently tight grid, we cross-checked these grid-searched parameters with estimates obtained by nonlinear least squares as well as those obtained through Feenstra's original methodology. Using our grid spacing, the difference between the parameters estimated using Feenstra's methodology and those obtained using the grid search differed only by a few percent for the 65 percent of sigmas for which we could apply Feenstra's approach.

One concern that one might have with this approach is that our grid search sets a maximum σ_g of 131.5. While this potentially could bias our results because we do not allow for infinite elasticities of substitution, in practice the bias is likely to be quite small. To see this, consider the following example. If σ_g is 3, a 50 percent decline in the lambda ratio (i.e., a doubling in level of varieties) corresponds to a 29 percent decline in the price index; if σ_g is 20, it corresponds to a 4 percent decline; and if σ_g is 131.5, a 0.5 percent decline. In other words, constraining the elasticity of substitution to lie below 131.5 will not cause us to identify substantial variety gains or losses when none are present.

For technical reasons, rather than performing the grid search over parameters (ω_g,σ_g), it was easier to grid search over (ρ_g,σ_g) where $\rho_g = \omega_g(\sigma_g - 1)/(1 + \sigma_g\omega_g)$ and is restricted to the following interval $0 \leq \rho_g < (\sigma_g - 1)/\sigma_g$ (see Feenstra [1994] for the derivation of this restriction).

changed by calculating the λ_g ratio for every good g (see equation (11)). Third, by combining our estimates of the elasticity of substitution with the measures of variety for each good, we obtain an estimate of how much the exact price index for good g moved as a result of the change variety growth. Finally, we can apply the ideal log weights to the price movements of each good in order to obtain an estimate of the movement of the aggregate price of imports using equation (12). Once we know how much import prices have changed, it is simple to apply equation (13) to calculate the welfare gain or loss from these price movements.

VI.A. Elasticities of Substitution

We now turn to our estimation of the elasticities of substitution. Given the tens of thousands of elasticities we estimate, it is impossible to report all of the results here. However, we can provide some sample statistics that can shed light on the plausibility of our estimates. There are three main priors that we have about these parameters. The first is that as we disaggregate, varieties are increasingly substitutable. In other words, to give a concrete example, varieties of the three-digit category of fruit and vegetables are likely to be less substitutable than varieties of the five-digit subcategory that only contains fresh, dried, or preserved apples. Similarly, varieties within this five-digit sector are likely to be still less substitutable than varieties in the seven-digit subcategory containing just fresh apples. Second, we would like the goods with high elasticities of substitution to correspond to goods that we think of as less differentiated. Finally, we would like to see that goods traded on organized exchanges have higher elasticities than those that are not.

Equation (20) can be estimated with g fixed at various levels of aggregation, and we report sample statistics for our elasticity estimates in Table IV.[22] The results reveal that for both time periods, as we disaggregate product categories, varieties appear to be closer substitutes. For instance, the simple average of the elasticities of substitution is 17 for seven-digit (TSUSA) goods during 1972–1988, while only 7 at the three-digit level. For the

22. A clarification can be handy to understand notation. When we estimate σ_g at the SITC-5 level, then c actually stands for the pair country-TSUSA goods. For instance, if two different TSUSA categories (e.g., Apples and Kiwis) belong to a given SITC-5 category (Fresh Fruit), then if the same country (Argentina) exports in the two TSUSA categories, the two pairs (Apples from Argentina and Kiwis from Argentina) will be treated as two different varieties of the same SITC-5 category (Fresh Fruit).

TABLE IV
SIGMAS FOR DIFFERENT AGGREGATION LEVELS AND TIME PERIODS

Period	Statistic	TSUSA/HTS	SITC-5	SITC-3
1972–1988	Mean*	17.3	7.5	6.8
	Standard error*	0.5	0.5	1.2
	Median	3.7	2.8	2.5
	Standard error	0.03	0.04	0.11
	Median varieties per category**	15	54	327
	Nobs of categories	11040	1457	246
1990–2001	Mean*	12.6	13.1 (6.6)	4.0
	Standard error*	0.5	5.9 (0.3)	0.5
	Median	3.1	2.7	2.2
	Standard error	0.04	0.06	0.13
	Median varieties per category**	18	52	664
	Nobs of categories	13972	2716 (2715)	256

(*) Estimates of the mean and standard error are adjusted for parameter censoring. The numbers in brackets in the SITC-5 1990–2001 were calculated dropping the one outlier elasticity of 16049.
(**) As in Table III, a variety is defined as a TSUSA/HTS-country pair.
For the TSUSA/HTS column: number of observations is equivalent to the median number of countries. For SITC-5 (SITC-3) column, it is the median number of TSUSA/HTS-good/country pairs in a given SITC-5 (SITC-3) level.

period between 1990 and 2001, the average elasticity was around 12 for ten-digit (HTS) goods and 4 among three-digit goods. These differences are not only large economically, but we can statistically reject the hypothesis that the mean coefficient for disaggregated goods is the same as that for more aggregated goods.[23] In terms of medians, the elasticity falls less dramatically, from 3.7 and 3.1 at the lowest levels of disaggregation in the first and second period, respectively, to 2.5 and 2.2. However, in both periods we can statistically reject that the medians at different levels of aggregation are the same. In sum, depending on the statistic being used, the elasticities of substitution fall by 33 to 67 percent at we move from highest to lowest level of disaggregation in Table IV. Note also that the median elasticities of substitution for a given disaggregation level tend to slightly fall over time and

23. We performed this test two ways. First, we tested the difference between the means of the estimated σ_g's, and second we recomputed the means and standard errors after accounting for the censoring of the σ_g's due to the grid search. In both cases, we can reject the hypothesis that the means are the same. We reported only the latter.

that these differences are statistically significant for the most and least disaggregated data. This finding is robust at all product levels, and may represent increasing differentiation among tradable goods in the latter period.[24]

Table V shows the elasticities of substitution for the twenty largest SITC-3 sectors in U. S. imports in each of the periods. For the period between 1972 and 1988, the sector with the highest elasticity of substitution among this group was that of crude oil. The estimated sigma for this sector was 17.1, fourteen times larger than the sigma for Footwear ($\sigma_{footwear} = 1.2$), the sector with the smallest elasticity in the table. In the latter period, we also find that sectors related to petroleum have the highest elasticities. More generally, a comparison of elasticities of substitution across categories shows an intuitive pattern that by and large seems reasonable.

Another way to establish the reasonableness of the estimates is to examine how well they correspond to other measures of homogeneous and differentiated goods. Rauch [1999] divided goods into three categories—commodities, reference priced goods, and differentiated goods—based on whether they were traded on organized exchanges, were listed as having a reference price, or could not be priced by either of these means. Commodities are probably correlated with more substitutable goods, but one should be cautious in interpreting commodities as perfect substitutes or the classification scheme as a strict ordering of the substitutability of goods. For example, although tea is classified by Rauch as a commodity, it is surely quite differentiated. Similarly, it is hard to see why a commodity like "dried, salted, or smoked fish" would be more homogeneous than a reference priced good like "fresh fish" or a differentiated good like "frozen fish." That said, it would be disturbing if we found that goods traded on exchanges are not more substitutable than those that are not.

In order to test this directly, we reestimated sigmas at the four-digit level to make them directly comparable with Rauch's classification and report the results in Table VI. The most striking feature of the table is that in both time periods, the average elasticities of substitution are much higher for commodities than

24. The total number of elasticities being estimated at the TSUSA/HTS level is smaller than the total number of TSUSA/HTS available within each period. This responds to the fact that the United States imports in a number of categories from a small number of countries, and we require at least three countries per category to identify parameters.

TABLE V
SIGMAS FOR THE TEN SITC-3 SECTORS WITH THE LARGEST IMPORT SHARE BY PERIOD

Period 1972–1988

SITC-3	Sigma	Average share (in %)	Descriptions
333	17.1	29.6	CRUDE OIL FROM PETROLEUM OR BITUMINOUS MINERALS
781	1.6	8.3	MOTOR CARS & OTHER MOTOR VEHICLES
334	9.0	5.4	OIL (NOT CRUDE) FROM PETROL & BITUMINOUS MINERALS, ETC.
341	5.7	2.4	LIQUIFIED PROPANE AND BUTANE
71	2.5	2.0	COFFEE AND COFFEE SUBSTITUTES
776	1.6	1.6	THERMIONIC, COLD CATHODE, PHOTOCATHODE VALVES, ETC.
641	6.7	1.6	PAPER AND PAPERBOARD
851	1.2	1.4	FOOTWEAR
681	1.4	1.2	SILVER, PLATINUM & OTHER PLATINUM GROUP METALS
674	11.8	1.2	IRON & NA STEEL FLAT-ROLLED PRODUCTS, CLAD, ETC.

Period 1990–2001

SITC-3	Sigma	Average share (in %)	Descriptions
781	3.0	10.6	MOTOR CARS & OTHER MOTOR VEHICLES
333	22.1	7.0	CRUDE OIL FROM PETROLEUM OR BITUMINOUS MINERALS
776	1.2	6.4	THERMIONIC, COLD CATHODE, PHOTOCATHODE VALVES, ETC.
752	2.2	5.7	AUTOMATIC DATA PROCESS MACHS AND UNITS THEREOF
784	2.8	3.4	PARTS AND ACCESSORIES OF MOTOR VEHICLES, ETC.
851	2.4	2.0	FOOTWEAR
764	1.3	1.9	TELECOMMUNICATIONS EQUIPMENT, N.E.S. AND PTS, N.E.S.
713	2.7	1.8	INTERNAL COMBUSTION PISTON ENGINES, AND PTS, N.E.S.
845	6.7	1.8	ARTICLES OF APPAREL OF TEXTILE FABRICS N.E.S.
641	2.1	1.7	PAPER AND PAPERBOARD

Shares are simple averages calculated over the entire period. Descriptions for SITC-3 codes are revision 3.

TABLE VI
ESTIMATED SIGMAS AND RAUCH LIBERAL CLASSIFICATION

	Rauch's classification of goods: Commodity	Reference priced	Differentiated
	1972–1988 (four-digit)		
Mean	15.3	7.8	5.2
Standard error	3.0	1.5	0.8
Test if different than commodity (p-value)		0.01	0.00
Median	4.8	3.4	2.5
Standard error	0.4	0.3	0.1
Test if different than commodity (p-value)		0.01	0.00
	1990–2001 (four-digit)		
Mean	11.6	4.9	4.7
Standard error	3.0	0.6	1.0
Test if different than commodity (p-value)		0.01	0.01
Median	3.5	2.9	2.1
Standard error	0.6	0.2	0.1
Test if different than commodity (p-value)		0.14	0.00

p-values for one-sided t-test reported.

for differentiated or reference priced goods, and the average elasticities of substitution for reference priced goods are higher than those of differentiated. The same picture emerges when we look at medians. In all but one case, we can strongly reject the hypothesis that commodities have the same average and median elasticity as reference priced goods and differentiated goods in both periods, and we can always reject the hypothesis that commodities have the same elasticity as the combined set of reference priced and differentiated goods. This suggests that goods that Rauch classifies as commodities are more likely to have high elasticities of substitution than goods that are classified as reference priced or differentiated.

VI.B. Growth in Varieties

Now that we have established that our estimates of the elasticities of substitution appear to be plausible by a number of criteria, we turn to the task of correctly evaluating changes in variety. One of the major obstacles we face in implementing this procedure is in the calculation of the λ_g ratio. Evaluating the impact on price of a new variety is straightforward to do in cases in which the United States imports other varieties of the same TSUSA/HTS category. Unfortunately, the λ_g ratio is undefined in cases where there are no common varieties of the TSUSA/HTS category between the start and end period (i.e., $I_g = \emptyset$ in Proposition 1). The reason why the λ_g ratio is undefined is that we cannot value the creation or destruction of a variety without knowing something about how this affects the consumption of other varieties. To give an example drawn from our data, we cannot value the invention of CD players for car radios without knowing how these new goods affected other goods, say, simple car radios. Our solution to this problem is to assume that whenever a new variety is created within a seven- or ten-digit category for which $I_g = \emptyset$ then all seven- or ten-digit categories within the same five-digit category have a common elasticity of substitution. In other words, in these special cases, the elasticity we use to evaluate the impact of a new variety being imported on the price level is a weighted average of the substitutability of other goods and varieties within the same five-digit category. Similarly, in cases where the entire five-digit category is new, we assume a common elasticity at the three-digit level.[25]

There are two important implications of this procedure for our results. The first is that the restriction on the set of goods for which we can calculate λ ratios means that for some product categories we need to define goods at the five-digit or three-digit level rather than at the TSUSA/HTS level. Because of this necessary restriction, instead of defining 12,347 goods in the earlier period and 14,549 goods for 1990–2001 (i.e., all TSUSA/HTS categories for which we have σ's),[26] we can only use 408 and 926 goods (a combination of TSUSA/HTS, SITC-5, and SITC-3), respectively. Note, however, that this forces some of the elasticities between different TSUSA/HTS categories to be the same, but the total number of varieties being used remains unchanged at over 150,000 and 250,000 in the period 1972–1988 and 1990–2001,

25. Note also that this approach also eliminates the bias arising from arbitrary recategorization of goods since new goods simply appear as new varieties of existing goods.

26. These are the numbers of available elasticities of substitution at the TSUSA and HTS level, respectively.

respectively. Moreover, we need to stress that this represents vastly more disaggregated data than has been used in the past. Whether this data limitation introduces a bias into our estimates is harder to assess.

Table VII shows descriptive statistics for the λ ratios of all the 1334 goods used in the calculation of the aggregate price index, and hence our sample statistics correspond to the complete set of imported varieties. As the table indicates, even when using λ ratios to measure variety growth, the typical sector saw the number of imported varieties increase. This table highlights the importance of using λ ratios rather than relying on count data to measure variety growth. As shown in Table I, the total number of varieties per TSUSA more than doubled in the period between 1972 and 1988 (i.e., $V_{72}/V_{88} = 0.46$). In turn, the number of HTS varieties rose by over 40 percent during 1990–2001. However, when we correctly account for the fact that varieties are not symmetric in the data, we find that the appropriate magnitudes of variety growth are substantially smaller. We find that the median measure of variety growth is approximately 25 percent (λ ratio = 0.81) in the period between 1972–1988 and 5 percent (λ ratio = 0.95) in the latter period. Although the count data suggest a 211 percent increase in the number of varieties, our λ ratios suggest that a 30 percent increase is more appropriate due to the large number of new varieties with small market shares. This underscores the importance of carefully measuring variety growth when making price and welfare calculations.

TABLE VII
DESCRIPTIVE STATISTIC OF LAMBDA RATIOS

Period	Statistic	Combination of TSUSA/HTS – SITC5 – SITC3 used	Implied by count data in Table I
1972–1988	Percentile 5	0.06	
	Median	0.81	0.46
	Percentile 95	2.00	
	Nobs	408	
1990–2001	Percentile 5	0.34	
	Median	0.95	0.70
	Percentile 95	1.80	
	Nobs	926	

See text for definitions.

TABLE VIII
THE IMPACT OF VARIETY IN U. S. IMPORT PRICES

Years	Ratio between aggregate exact price index including variety and the aggregate conventional price index	
	End-point ratio	Average per-annum ratio
1972–1988	0.803 [0.790,0.835]	0.986
1990–2001	0.917 [0.907,0.941]	0.992
1972–2001 (*)	0.720 [0.705,0.771]	0.988

This table shows the estimated values of equation (12) by period. (*) For the period between 1988 and 1990 the average per-annum rate was applied. Bootstrapped ninetieth percent confidence intervals are in brackets.

VI.C. Import Prices and Welfare

We are now ready to use the elasticities of substitution to evaluate the price effects of changes in varieties. Aggregating together our λ ratios according to equation (12) yields estimates of the impact of variety growth on the exact aggregate import price index. The results from this exercise are reported in Table VIII. Standard errors on the bias were computed by bootstrapping each estimate of σ_g 50 times and recalculating the bias for each set of parameters. Overall, variety growth implies that the variety adjusted unit price for imports fell a precisely estimated 19.7 percent faster than the unadjusted price between 1972 and 1988 or about 1.4 percentage points per year. Interestingly, the impact of variety growth was much smaller during the 1990s. Between 1990 and 2001 the growth of varieties meant that the exact price index fell 8.3 percent faster than the unadjusted index over this time period or about 0.8 percentage points per annum. The lower rate of decline in the later period may reflect the fact that much of the gains from globalization arising from rise in importance of East Asian trade may have been realized prior to 1990. If we assume that prices declined in the missing year at the average rate across the entire sample, we find that throughout the entire period, the growth of varieties reduces the exact price relative to conventionally measured import price index by 28.0 percent.

It is difficult to find a benchmark with which to compare our results. We are not aware of any study that measures the impact of variety on aggregate prices, and the papers that study a single good at the micro-level (or at most a few goods) are not suitable for this comparison. Given the lack of aggregate effects of variety in the literature, we will use as a reference the effects that other sources of bias (quality change, outlet substitution, etc.) have on the overall consumer price index. In mid-1995 a commission was appointed to study the potential biases in the existing measurement of the Consumer Price Index. This CPI Commission concluded that the change in the consumer price index overstates the change in the cost of living by about 1.2 percentage points per year [Boskin et al. 1996]. Several sources of bias are considered, but the main source is the incorrect measurement of quality change of products. The effect of quality change alone can account for about 0.6 percentage points in the overall index. These numbers suggest that the bias that we find in the import price index only as a result of the unaccounted variety growth is very large. That is, the bias due to variety growth in the import price index is almost twice as large as the bias induced by quality change in the overall price index and as large as the total bias from all sources.

We now turn to calculating the welfare effect of the fall in the U. S. exact import price. Not surprisingly, the magnitude of the welfare gain from this fall hinges on the functional forms underlying the Dixit-Stiglitz structure and cannot be general (an issue we will return to in the next section). If elasticities of substitution are not constant or if marginal costs are not fixed, theory suggests that one can obtain higher or lower estimates of the gains from variety. Although our estimate of the impact of imported varieties on import prices is correct for any domestic production structure, we cannot translate this into a welfare gain without making explicit assumptions about the structure of domestic production. Our choice is to assume the same structure of the U. S. economy as in Krugman [1980]. We do this for two reasons. First, since this is the dominant model of varieties, it provides a useful benchmark for understanding the potential welfare gains. Second, we lack the necessary data and model of the economy's input-output linkages to estimate variants of the monopolistic competition model in which there are more complex interactions between imported and domestic varieties.

The compensating variation that results from changes in varieties in imported goods can be calculated using the inverse of the product of the weighted λ ratios raised to the fraction of imported goods in total consumption goods, as shown in equation

(12). In particular, given equation (13), we use the ideal import share in each period, 6.7 percent for 1972–1988 and 10.3 percent for 1990–2001, respectively, together with the information in Table VIII to obtain the gains in welfare due to variety.[27] We find that consumers are willing to pay 2.6 percent of their income to access the wider set of varieties available in 2001 rather than the set available in 1972. Around 1.8 percentage points accrue to the earlier period. On a per-year basis, consumers are willing to pay on average 0.1 percent of their income to access each year's new set of varieties rather than staying with the set of the previous year.

VI.D. Robustness of Results to Alternative Assumptions

In the previous subsection we computed the impact of variety growth of U. S. imports on aggregate welfare. Our computation required several weaker assumptions than those present in previous numerical exercises. First, we do not require that varieties or goods have equal shares in consumption. Second, we allow for different elasticities of substitution for each of the goods used. Third, we obtain our elasticities of substitution by allowing each of our 1334 markets to have a different elasticity of supply rather than assuming that these supply elasticities are always equal to zero. Table IX underscores the importance of using this weaker set of assumptions. As mentioned in Section V, when import shares are assumed equal for all varieties, the aggregate import bias in equation (12) becomes

$$\prod_{g \in G} \left(\frac{V_{gt-1}}{V_{gt}} \right)^{w_{gt}(G)/(\sigma_g - 1)},$$

where V_{gt-1}/V_{gt} is the ratio of the actual number of varieties in each period. Column (2) of Table IX shows how the impact of variety on welfare is affected by using a simple count of varieties (i.e., the V ratios) rather than the appropriate λ ratios. By using V ratios, the welfare gains from variety growth become 6.28 percent, more than twice as big as the true estimate using λ ratios. This suggests that in the case of U. S. imports, using a simple count of varieties to measure the impact of variety growth grossly overestimates the true impact.

27. We assume that welfare rose at the geometric average of the rates in the two periods between 1988 and 1990.

TABLE IX
WELFARE COMPARISONS

	Benchmark estimate (1)	Equality of import shares (2)	Single gravity sigma (3)	Smaller median sigma (4)
Number of sigmas used	1334	1334	1	1
Median sigma	2.9	2.9	2.9	2.0
Median [percentile 5,percentile 95]	[2.1,5.2]			
Average sigma	6.5	6.5	2.9	2.0
Lambda ratio (L) or count data (N)	L	N	N	N
Average share (in percent) (*)	8.5	8.5	8.5	8.5
Welfare impact	2.59 [2.01,2.84]	6.28	5.09	8.62

(*) For expositional purposes, we present the weighted average share of imports between the periods 1972–1988 and 1990–2001. Calculations are based on each period's average share and not on this weighted average.

Column (3) shows the additional impact of using a single elasticity of substitution for all varieties. For comparison purposes with column (2), we choose our median elasticity 2.9 as a benchmark. The impact of variety on welfare using the single elasticity of 2.9 is 5.09 percent, smaller than that of column (2) but still almost twice as large as our benchmark estimate. This underscores the importance of using the full distribution of sigmas to value variety. Despite using a smaller point estimate of the average elasticity (relative to columns (1) and (2)), we find that the impact of variety on welfare is reduced.[28] Finally, column (4) shows the effect of a reduction in the average elasticity of substitution on welfare. By reducing the average from 2.9 to 2.0 (as used in Romer [1994]), the impact on welfare significantly rises to 8.62 percent. In sum, columns (2)–(4) quantify the importance of using λ ratios and a complete distribution of sigmas to calculate the impact of variety growth on welfare.

The previous discussion illustrates that within the CES framework our estimated σ's and λ ratios tend to result in *smaller* price and welfare movements than one would obtain if count data

28. This reveals that sectors with high variety growth seem to be associated with low elasticity of substitutions.

or a common elasticity were used. However, what happens if we relax the maintained assumption of a CES utility function? We can address this by recomputing our main results using an exact aggregate price index based on a quadratic utility function instead.

The benefit of using a quadratic utility function is that its exact price index is the Fisher ideal price index, which is the geometric average of the Paasche and Laspeyres indices. Specifically, we assume that utility can be written as

$$U = (D^{\psi} + (\mathbf{m}'\mathbf{A}\mathbf{m})^{\psi/2})^{1/\psi},$$

where $\mathbf{m}$ is column vector whose elements are the imported quantities of varieties (including varieties whose quantity is zero), ψ is a parameter between zero and one, and $\mathbf{A}$ is a symmetric matrix indicating how the varieties enter into utility. The import price index will then be the Fisher ideal price index given by

$$(21)\quad p^{M}(\mathbf{p}^{0},\mathbf{p}^{1},\mathbf{m}^{0},\mathbf{m}^{1}) \equiv [\mathbf{p}^{1'}\mathbf{m}^{1} \times \mathbf{p}^{0'}\mathbf{m}^{1}/(\mathbf{p}^{1'}\mathbf{m}^{0} \times \mathbf{p}^{0'}\mathbf{m}^{0})]^{1/2},$$

where the superscript zeros and ones correspond to the initial and final periods, respectively. The exact price index will then be a weighted average of the domestic and imported goods price indexes, where we rely on Sato and Vartia for the formulas for the weights.

The only remaining issue is how to measure the reservation prices of varieties. To do this, we use the fact that our elasticity of substitution corresponds to our estimate of the local demand elasticity or

$$(22)\qquad \frac{dM_{gc}}{dp_{gc}}\frac{p_{gct}}{M_{gcr}} = -\sigma_g.$$

If we assume that each variety constitutes a small component of consumption so that we can ignore the effect of the consumption of one variety on the aggregate price index, then the demand curve will be linear in the price of any variety. Without loss of generality, for a variety with positive imports in period t but not in period r, we can express the implied price in period r as

$$(23)\qquad M_{gct} = \frac{\partial M_{gc}}{\partial p_{gc}}\,(p_{gct} - p_{gcr}) \text{ or } p_{gcr} = p_{gct}\left(1 + \frac{1}{\sigma_g}\right),$$

where p_{gcr} corresponds to the reservation price of the variety.

TABLE X
IMPACT OF VARIETY CHANGE UNDER ALTERNATIVE ASSUMPTIONS

Price bias (1972–2001) with exact price indices (*)	
CES utility function	0.72
Quadratic utility (Fisher ideal price index)	0.67
Welfare gain (1972–2001) as a percent of GDP	
Utility-based approaches	
Benchmark CES utility	2.59
Quadratic utility	2.65

(*) Price biases are expressed as ratios between aggregate exact price index and the conventional price index (as in Table VI).

Using equation (23), we can compute the reservation prices for all appearing and disappearing goods and then plug these into equation (21).[29] By taking the ratio of the Fisher ideal price index computed using all varieties (i.e., including the new and disappearing varieties) relative to the index computed using only common goods, we obtain an estimate of the gain or loss in utility arising from prices of disappearing and created varieties moving to and from their reservation levels.

The result from this exercise is reported in Table X. The Fisher ideal price index that incorporates the virtual price movements for the disappearing and created goods is 33 percent lower than the Fisher index that is based solely on the common goods. Using the Sato-Vartia weights, we find that consumers would be willing to pay 2.65 percent of GDP for the expanded set of varieties. These numbers (33 percent and 2.65 percent) are startlingly close to the numbers we obtain with the Feenstra price index (28 percent and 2.59 percent). In other words, shifting between CES preferences and quadratic preferences that imply a finite reservation price has a very small impact on our results.

Finally, we assess the sensitivity of our estimates to two key assumptions of our framework. First, in our analysis all imported goods are assumed to be for final consumption. This is not the case in the data, as almost two-thirds of imports are intermediate

29. We wish to make two clarifying points about this calculation. First, we are computing the reservation prices assuming that the sector is small and all of the other prices are fixed. We therefore are not taking into account how the reservation prices might change when the prices of other goods move. Doing so would substantially complicate the calculation. Second, in order not to undervalue the welfare losses in the earlier time period, we first converted historical prices into current prices using the import price deflator before doing the calculation.

or capital goods and not final consumption goods. As noted by Romer [1994], however, the Dixit-Stiglitz structure allows for new varieties to be modeled either as consumption goods, as in Grossman and Helpman [1991], or as intermediate inputs in production, as in Romer [1990], with no fundamental change in the underlying economic analysis. In other words, treating the share of imported intermediate goods as final consumption does not necessarily bias our estimates. The case of capital goods is different. While consumption varieties only offer a static gain, the potential gains from variety growth in capital goods can have persistent effects in time. This has the potential of magnifying the effects of variety growth on welfare relative to our "static" results. Therefore, treating capital goods as final consumption seems to imply that we are understating the true gains from variety.

The second assumption that is important for our results is that of Krugman's [1980] production structure. This structure assumes that the number of domestic varieties is unaffected by new foreign varieties (as in Feenstra [1992], Romer [1994], and Klenow and Rodriguez-Clare [1997]). However, if domestic varieties were affected by imported varieties, then our welfare calculations would change. One way to assess the impact of this change is to calibrate the model of Helpman and Krugman [1985, pp. 197–209]. This model of the monopolistic competition framework postulates that each country produces both traded and nontraded goods. The entry of foreign varieties will cause the domestic differentiated sector to shrink as firms move into the nontraded sector. In the interest of brevity, we will not rederive the key equations. In this model, the demand of foreign goods relative to domestic goods can be written as

$$n^*D^f/nD = n^*\tau^{-\sigma}/n, \tag{24}$$

where n and n^* are the number of domestic and foreign firms, D and D^f are the quantity of each domestic and foreign variety purchased by the home country, $\tau > 1$ is the iceberg transportation cost, and σ is the elasticity of substitution. One can also show that if one holds fixed the domestic labor force and assumes both countries produce differentiated goods, then

$$dn = -\tau^{1-\sigma}dn^*. \tag{25}$$

If we assume that the left-hand side of equation (24) equals the share of imports to domestic demand (= GDP − exports) in 2001 (i.e., 0.13), n^*/n equals the ratio of GDP in the rest of the world to U. S. GDP (i.e., 2.13), and σ equals our median sigma of 2.9, then after a little algebra one can show that $-\tau^{1-\sigma} = -0.16$. This suggests that every incoming variety displaces 0.16 domestic varieties, and hence one might want to reduce our welfare estimate by 16 percent, i.e., from 2.6 percent of GDP to 2.2 percent of GDP. In this context, making the number of domestic firms endogenous reduces our point estimate, but it does not dramatically alter our main result.

VII. Conclusion

Understanding the impact of new products and the growth in varieties on economies has been one of the central questions in international economics, regional economics, and macroeconomics. Until now, attempts to estimate magnitude of these effects have been limited to extremely careful econometric studies of particular goods and attempts to calibrate standard models. The failure to obtain credible estimates of the impact of new goods and varieties on prices and welfare at the national level has stemmed from the difficulty of implementing careful econometric studies of particular markets more broadly and from the implausibility of many assumptions underlying calibration exercises.

This paper provides a methodology that estimates all of the parameters necessary to calculate an exact aggregate price index and perform welfare calculations at the national level. This, of course, does not obviate the need for careful econometric studies of entry in particular markets. Indeed, we see this work as complementary to ours. In markets where sufficient data exist to obtain better estimates of price effects due to entry, one should do so. Indeed, one could imagine more precise estimates of the impact of new goods and varieties arising from a hybrid technique in which certain markets are modeled in detail and others are modeled according to our implementation of the Feenstra index. Whether that would substantially alter our results is impossible to say, but our results are robust to a wide range of model specifications and assumptions.

Our results indicate that the effect of new goods and varieties on the U. S. economy is large. By ignoring these effects, a conventional import price index overstates import price inflation by 1.2 percentage points per year. If the bias in the CPI is comparable to the bias in the import price index, then this suggest that the biases identified in this paper are not only more important than quality adjustments, they are more important than all other biases *combined.* Obviously, more work needs to be done to understand whether the biases in the import price index are similar in magnitude to those in the CPI, but the results suggest that there is potential for very large effects.

Finally, our results suggest that globalization had had substantial impacts on welfare through the import of new varieties. U. S. welfare is 2.6 percent higher due to gains accruing from the import of new varieties. An important qualification is that our estimates are obtained by assuming the U. S. economy can be modeled as in Krugman [1980]. While this is a sensible benchmark, there clearly is a need for better modeling and estimation of dynamic and input-output effects arising from increases in the number of imported varieties. Even so, our estimation indicates that the gains from trade first suggested by Krugman a generation ago are quite important in reality.

Appendix: Biases and Weighting in the Presence of Measurement Error

Let p_{gct} be the unit value of a variety that we have in our data set (i.e., French red wine) and p_{gcti} is the price of product i contained in variety gc (i.e., the price of a particular bottle of French red wine). Hence q_{gcti} will always equal one. If we assume that the log of the geometric mean price of a variety is approximately equal to the log of the arithmetic mean, we have

$$\ln p_{gct} \equiv \ln\left(\frac{\sum_i p_{gcti}}{q_{gct}}\right) \approx \ln\left(\left(\prod_i p_{gcti}\right)^{1/q_{gct}}\right), \tag{26}$$

where $q_{gct} \equiv \sum_i q_{gcti}$ is the quantity of imported variety gc in time t.

Assume that product prices are measured with an i.i.d. error such that $p_{gcti} = \tilde{p}_{gcti}\zeta_{gcti}$, where $\tilde{p}_{gcti}$ is the true price and p_{gcti} is the measured price. In this case

$$\mathrm{var}(\ln \zeta_{gcti}) \equiv \chi^2$$
$$\mathrm{cov}(\ln \zeta_{gcti}, \ln \zeta_{gc'sj}) = 0 \;\; \forall\, c \neq c', t \neq s, i \neq j$$

$$(27)\quad \chi^2_{\ln p_{gct}} \equiv \operatorname{var}\left[\ln\left(\frac{\sum_i p_{gcti}}{q_{gct}}\right)\right] \approx \operatorname{var}\left[\ln\left(\left(\prod_i p_{gcti}\right)^{1/q_{gct}}\right)\right]$$
$$= \frac{1}{q_{gct}^2}\operatorname{var}\left(\sum_i \ln p_{gcti}\right) = \frac{1}{q_{gct}^2} q_{gct}\chi^2 = \frac{1}{q_{gct}}\chi^2.$$

Now

$$(28)\qquad E(\ln p_{gct} - \ln p_{gct-1})^2 = \delta^2_{gct} + \chi^2\left(\frac{1}{q_{gct}} + \frac{1}{q_{gct-1}}\right),$$

where δ^2_{gct} is the variance of the true price change. Averaging this across all periods produces

(29)

$$E\,\frac{1}{T}\sum_t (\ln p_{gct} - \ln p_{gct-1})^2 = \frac{1}{T}\sum_t \delta^2_{gct} + \chi^2\,\frac{1}{T}\sum_t\left(\frac{1}{q_{gct}} + \frac{1}{q_{gct-1}}\right).$$

This implies that we should add a term equal to

$$(30)\qquad \widehat{\chi^2}\,\frac{1}{T}\sum_t\left(\frac{1}{q_{gct}} + \frac{1}{q_{gct-1}}\right)$$

to the right-hand side of equation (18) where $\widehat{\chi^2}$ is a parameter to be estimated. It is worth noting that if $q_{gct} = q_{gct-1}$ and the importance of measurement error does not decline with the number of periods used in the between estimate, the terms in parentheses in equation (30) become a constant. This is the specification estimated in Feenstra [1994]. Our algebra can thus be seen as a generalization of Feenstra's approach that allows for measurement error to depend on the quantity of varieties and the number of periods over which the average is taken.

A related but distinct issue concerns heteroskedasticity in the data. Every data point comprising the left-hand side of equation (18) is an estimate of the variance. However, if the prices are measured with error, then so are our sample variances. Using equation (28), we can correct for this heteroskedasticity by realizing that the variance of the average of the sample variance can be written as

$$(31)\quad \operatorname{var}\left\{\frac{1}{T}\sum_t\left[(\ln p_{gct}-\ln p_{gct-1})^2-\frac{1}{T}\sum_t(\ln p_{gct}-\ln p_{gct-1})^2\right]^2\right\}$$
$$=\frac{1}{T^2}\operatorname{var}\left\{\sum_t\left[(\ln p_{gct}-\ln p_{gct-1})^2-\frac{1}{T}\sum_t(\ln p_{gct}-\ln p_{gct-1})^2\right]^2\right\}.$$

The term in curly brackets is likely to be inversely related to the quantity of goods used in order to compute the unit value and inversely related to the number of periods over which we compute the average variance. We correct for this heteroskedasticity by assuming that the variance of the average sample variance is proportional to

$$\frac{1}{T^3}\left(\frac{1}{q_{gct}}+\frac{1}{q_{gct-1}}\right),$$

and we therefore weight the data by

$$T^{3/2}\left(\frac{1}{q_{gct}}+\frac{1}{q_{gct-1}}\right)^{-1/2}.$$

UNIVERSITY OF CHICAGO, GRADUATE SCHOOL OF BUSINESS, AND FEDERAL RESERVE BANK OF NEW YORK
COLUMBIA UNIVERSITY AND NATIONAL BUREAU OF ECONOMIC RESEARCH

REFERENCES

Anderson, James, "A Theoretical Foundation for the Gravity Equation," *American Economic Review,* LXIX (1979), 106–116.
Armington, Paul, "A Theory of Demand for Products Distinguished by Place of Production," *International Monetary Fund Staff Papers,* XVI (1969), 159–178.
Bils, Mark, and Peter Klenow, "The Acceleration of Variety Growth," *American Economic Review,* XCI (2001), 274–280.
Boskin, Michael, Ellen Dulberger, Robert Gordon, Zvi Griliches, and Dale Jorgenson, *Toward a More Accurate Measure of the Cost of Living: Final Report to the Senate Finance Committee* (Washington, DC: U. S. Government Printing Office for the U. S. Senate Committee on Finance, December 1996).
Broda, Christian, and David Weinstein, "Variety Growth and World Welfare," *American Economic Review Papers and Proceedings,* XCIV (2004), 139–145.
Brown, Drusilla, Alan V. Deardorff, and Robert Stern, "Modeling Multilateral Trade Liberalization in Services," *Asia-Pacific Economic Review*, II (1996), 21–34.
Diewert, W. Erwin, "Exact and Superlative Index Numbers," *Journal of Econometrics,* IV (1976), 115–145.
Dixit, Avinash, and Joseph Stiglitz, "Monopolistic Competition and Optimum Product Diversity," *American Economic Review,* LXVII (1977), 97–308.
Eaton, Jonathan, and Samuel Kortum, "Technology, Geography, and Trade," *Econometrica,* LXX (2002), 1741–1779.
Eaton, Jonathan, Samuel Kortum, and Francis Kramarz, "Dissecting Trade: Firms, Industries, and Export Destinations," *American Economic Review Papers and Proceedings,* XCIV (May 2004), 150–154.

Feenstra, Robert, "How Costly Is Protectionism?" *Journal of Economic Perspectives,* VI (1992), 159–178.
——, "New Product Varieties and the Measurement of International Prices," *American Economic Review,* LXXXIV (1994), 157–177.
——, "U. S. Imports, 1972–1994: Data and Concordances," National Bureau of Economic Research Working Paper No. 5515, 1996.
Feenstra, Robert, Gordon Hanson, and Songhua Lin, "The Value of Information in International Trade: Gains to Outsourcing through Hong Kong," *Advances in Economic Analysis & Policy,* IV (2004), 1–35.
Fisher, Irving, *The Making of Index Numbers* (Boston, MA: Houghton-Mifflin, 1922).
Fujita, Masahisa, Paul Krugman, and Anthony Venables, *The Spatial Economy: Cities, Regions, and International Trade* (Cambridge, MA, and London, UK: MIT Press, 1999).
Grossman, Gene, and Elhanan Helpman, *Innovation and Growth in the Global Economy* (Cambridge, MA, and London, UK: MIT Press, 1991).
Hansen, Lars, "Large Sample Properties of Generalized Method of Moments Estimators," *Econometrica,* L (1982), 1029–1054.
Hausman, Jerry, "Exact Consumer's Surplus and Deadweight Loss," *American Economic Review,* LXXI (1981), 662–676.
Helpman, Elhanan, and Paul Krugman, *Market Structure and Foreign Trade: Increasing Returns, Imperfect Competition, and the International Economy* (Cambridge, MA, and London, UK: MIT Press, 1985).
Hotelling, H., "Stability in Competition," *Economic Journal,* XXXIX (1929), 41–57.
Kemp, Murray, "Errors of Measurement and Bias in Estimates of Import Demand Parameters," *Economic Record,* XXXVIII (1962), 369–372.
Klenow, Peter, and Andres Rodriguez-Clare, "Quantifying Variety Gains from Trade Liberalization," September 1997.
Krugman, Paul, "Increasing Returns, Monopolistic Competition, and International Trade," *Journal of International Economics,* IX (1979), 469–479.
——, "Scale Economies, Product Differentiation, and the Pattern of Trade," *American Economic Review,* LXX (1980), 950–959.
Lancaster, Kelvin, "Socially Optimal Product Differentiation," *American Economic Review,* LXV (1975), 567–585.
Rauch, James, "Networks versus Markets in International Trade," *Journal of International Economics,* XLVIII (1999), 7–35.
Rigobon, Roberto, "Identification Through Heteroskedasticity," *Review of Economics and Statistics,* LXXXI (2003), 777–792.
Roberts, Mark, and James R. Tybout, "The Decision to Export in Colombia: An Empirical Model of Entry with Sunk Costs," *American Economic Review,* LXXXVII (1997), 545–563.
Romer, Paul, "Capital, Labor, and Productivity," *Brookings Papers on Economic Activity, Special Issue* (1990), 337–367.
——, "New Goods, Old Theory, and the Welfare Costs of Trade Restrictions," *Journal of Development Economics,* XLIII (1994), 5–38.
Rutherford, Thomas, and David Tarr, "Trade Liberalization, Product Variety and Growth in a Small Open Economy: A Quantitative Assessment," *Journal of International Economics,* LVI (2002), 247–272.
Samuelson, Paul, "Using Full Duality to Show that Simultaneously Additive Direct and Indirect Utilities Implies Unitary Price Elasticity of Demand," *Econometrica,* XXXIII (1965), 781–796.
Sato, Kazuo, "The Ideal Log-Change Index Number," *Review of Economics and Statistics,* LVIII (1976), 223–228.
Spence, Michael, "Product Differentiation and Welfare," *American Economic Review,* LXVI (1976), 407–414.
Vartia, Yrjo, "Ideal Log-Change Index Numbers," *Scandinavian Journal of Statistics,* III (1976), 121–126.
Yi, Kei-Mu, "Can Vertical Specialization Explain the Growth of World Trade?" *Journal of Political Economy,* CXI (2003), 52–102.