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ABSTRACT

This paper is the first attempt to structurally estimate the impact of globalization on markups and welfare in a monopolistic competition model. To achieve this, we work with a class of preferences that allow for endogenous markups and firm entry and exit that are especially convenient for empirical work – the translog preferences, with symmetry in substitution imposed across products. Between 1992 and 2005 we find the U.S. market experienced a series of changes that confirm the predictions of Melitz and Ottaviano (2008): import shares rose and U.S. firms exited, leading to a fall in markups, while product variety and welfare went up. We estimate the impacts of these effects on a national level, and find a cumulative drop of 5.4 percent in merchandise prices and of 1.0 percent in overall consumer prices between 1992 and 2005. Although the magnitude of the welfare gains in our translog setup is similar to that obtained by assuming CES preferences, the sources of these gains are quite different. Variety gains under translog are at least one-third smaller than in the CES case, but there is a substantial reduction in U.S. markups, resulting in a comparable welfare gain overall.

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1. Introduction

A promise of the monopolistic competition model in trade was that it offered additional sources of the gains from trade, beyond that from comparative advantage (e.g. Krugman (1979) and more recently Melitz and Ottaviano (2008)). These additional sources include: consumer gains due to the expansion of import varieties; efficiency gains due to increasing returns to scale; and welfare gains due to reduced markups. While the first two sources of gains have received recent empirical attention,[1] the promise of the third source – reduced markups – has not yet been realized. To be sure, there are estimates of reduced markups due to trade for several countries: Levinsohn (1993) for Turkey; Harrison (1994) for the Ivory Coast; and Badinger (2007a) for European countries. But these cases rely on dramatic liberalizations to identify the change in markups and are not tied in theory to the monopolistic competition model. The reason that this model is not used to estimate the change in markups is because of the prominence of the constant elasticity of demand (CES) system, with its implied constant markups. To avoid that case, the above authors do not specify the functional form for demand and instead rely on a natural experiment to identify the change in markups.

For these reasons, we do not have evidence beyond these case studies about how the broad process of globalization affects markups, and particularly no evidence on the impact of such markup reductions on U.S. welfare. This paper is the first attempt to structurally estimate the impact of globalization on markups and welfare in a monopolistic competition model. To achieve that, we work with a class of preferences that are new to that literature – the translog preferences, with symmetry in substitution imposed across products. These preferences are

[1] The consumer gains due import variety have been estimated for the U.S. by Broda and Weinstein (2006). Gains due to increasing returns to scale, or more specifically due to the self-selection of efficient firms (as in Melitz, 2003) have been demonstrated for Canada by Trefler (2004) and for a broader sample of countries by Badinger (2007b, 2008). See also Head and Ries (1999, 2001) for Canada, and Tybout *et al* (1991, 1995) for Chile and Mexico.

known to have good properties for empirical work (Diewert, 1976): they are homothetic; can give a second-order approximation to an arbitrary expenditure function; and correspond to the Törnqvist price index, which is very close to price index formulas that are used in practice. Furthermore, these preferences prove to be highly tractable even as the range of import varieties change, so they can join the quadratic preferences used by Melitz and Ottaviano (2008) as being alternatives to the CES case that allow for endogenous markups.[2]

In the translog case the elasticity of demand is inversely related to a product's market share, so markups fall as more firms enter, which we call the *pro-competitive effect*. On the other hand, domestic firms may exit as foreign competition intensifies, offsetting some of this gain to consumers. This we will refer to as the *domestic exit effect*. Incorporating these two effects into the analysis allows us to estimate the impact of globalization on markups. Furthermore, this class of preferences also allows us to address a potential criticism of Broda and Weinstein (2006): that by assuming CES preferences, it may overstate the gains from import variety.[3] The translog system allows for an alternative estimate of the variety gains, which we find are at least one-third smaller than in the CES case. But our *combined* gains for the U.S. due to import variety and the pro-competitive effect are of the same magnitude as Broda and Weinstein's CES estimates.

Our results are broadly consistent with the predictions of Melitz and Ottaviano (2008) in that the globalization of the U.S. economy between 1992 and 2005, as measured by the rise in import penetration, was associated with a substantial decline in the number of producers in the

[2] The quadratic preferences used by Melitz and Ottaviano (2008) lead to linear demand curves with zero income elasticity, though country population can act as a demand shift parameter. Demand curves of this type and the associated markups are estimated in the industrial organization literature: see Bresnahan (1989) and the recent trade application by Blonigen *et al* (2007). For other preferences that are non-homothetic and allow for variable markups see Behrens *et al* (2008) and Simonovska (2008).

[3] The gains from a new product variety can be thought of as the area under the demand curve and above the price when the product first appears. While the CES system has an infinite reservation price, this area under the demand curve is still bounded above (provided the elasticity of substitution is greater than unity). But it can be expected that the gains from new product varieties in this case might exceed the gain from other functional forms with *finite* reservation prices, as is the case for the translog system.

U.S. As a result, U.S. Herfindahl indexes rose in many markets. The product of the Herfindahl index and the overall U.S. market share measures the market share of the typical U.S. firm. While the Herfindahls rose due to exit, the overall U.S. market share fell by more, so it follows that the typical *per-firm* share of a surviving U.S. producer also fell. That finding provides us with *prima facie* evidence that there has been an increase in competition and reduced markups. In fact, for the translog system, the sum of the Herfindahl indexes for U.S. producers and for exporters to the U.S., weighted by their squared market shares, is precisely the right way to measure competition, and we show that these “market-level” Herfindahl indexes have fallen in many sectors.

Our results suggest that globalization has been exerting important economic impacts on the U.S. economy. Our point estimate for the cumulative gains to U.S. consumers from new varieties and decreased markups is 1.0 percent over the period 1992 to 2005. However, the impact on the merchandise sector (agriculture, manufacturing, and mining) was much larger. The welfare gains in that sector was equivalent to a 5.4 drop in prices, with 1.7 percentage points coming from reduced markups and 3.7 percentage points from new varieties.

In section 3, we introduce the translog expenditure function and solve for the ratio of expenditure functions (or exact price index) in the presence of new and disappearing goods, which allow the gains from new products to be measured. The pro-competitive effect of imports is discussed in section 4. Our analysis allows for multiple products supplied by each country, and shows how the Herfindahl indexes of export sales by each country enter into our equations. Significantly, we have been able to obtain these indexes for most countries selling to the U.S., by land or by sea. In section 5 we discuss the procedure for estimating the system of demand and pricing equations, and results are presented in section 6.

2. Data Preview

One of the dramatic changes that globalization has wrought on the U.S. economy is the declining share of U.S. demand supplied by plants located in the U.S. To see this, we define U.S. domestic supply as aggregate U.S. sales less exports for agricultural, mining, and manufacturing goods (see the data Appendix for detailed definitions of all of our variables). We define U.S. apparent consumption as domestic supply plus imports. Similarly, we define the U.S. suppliers' share of the U.S. market, as U.S. domestic supply divided by apparent consumption. Finally, we define each country's U.S. import share as the exports from that country to the U.S. divided by apparent consumption.

The switch in U.S. classification of output data from the SIC system to the NAICS in 1998 makes it difficult to compare sectoral output levels between 1997 and 1998. We therefore break our sample into two periods (1992-1997 and 1998-2005) to maintain consistent series, and discuss how we handle this problem for the estimation in the appendix. For the initial tables, we will present the raw data drawn from two subsamples, but we will present results for both sub-samples and the full sample in the results section.

From Table 1 we see that the share of U.S. apparent consumption sourced domestically fell by a little more than 5 percentage points between 1992 and 1997 and by 9 percentage points between 1998 and 2005. This decline corresponds to an annual decline in the U.S. share of 1.4 percentage points per year in the early period and 1.7 percentage points per year in the later period. The flip side of this decline was an almost doubling of the import share. Interestingly, the growth of imports was not uniform across countries: depending on the time period, between one-half and two-thirds of the increase was due to increases in import shares from Canada, China, and Mexico – countries that were either growing rapidly or involved in free trade agreements.

One possible explanation for the findings in Table 1 is that the rise in import penetration was confined to a few important sectors. We can examine whether this was the case by looking at more disaggregated data. In Figure 1, we plot the U.S. suppliers' share in 1997 or 2005 against its level in 1992 or 1998, for each HS 4-digit category. We also place a 45-degree line in the plot so that one can easily see which sectors experienced gains in U.S. shares and which experienced declines. As one can see from the figure, the vast majority of sectors lie below the 45-degree line, meaning that import penetration was steadily expanding over this time period. This establishes that the rise in import penetration, though quite pronounced in some sectors, was a general phenomenon that was common across many merchandise sectors.

Along with the declining U.S. market share in many sectors, there has also been an exit of manufacturing firms. The Department of Census data reveals that in 1992, there were 337,409 firms in manufacturing. By 2002 this number had fallen to 309,696: an 8.2 percent decline. We will argue that this decline in the number of firms was also associated with an even larger decline in U.S. market share, resulting in not only a rise in imports but also a decline in the typical market share of a surviving U.S. firm. Thus, by 2005 the U.S. market was characterized by fewer domestic firms with smaller per-firm shares.

To make this clear, it is convenient to work with Herfindahl indexes of market concentration, defined for each country selling to the U.S. We let i denote countries, j denote firms (each selling one product), k denote sectors and t denote time. Let s_{jt}^{ik} denote firm j's exports to the U.S. in sector k, as a share of country i's total exports to the U.S. in that sector. Then the Herfindahl for country i is:

$$H_{it}^{k} = \sum_{j} \left(s_{jt}^{ik} \right)^{2} \ .$$

The inverse of a Herfindahl can be thought of as the "effective number" of exporters, or U.S. firms, in an industry. Thus, a Herfindahl of one implies that there is one firm in the industry and an index of 0.5 would arise if there were two equally sized firms in the sector. Similarly, if we multiply the Herfindahl by the share of the country's suppliers in the market, one obtains the market share of a synthetic typical firm in the market. This is a very useful statistic because in many demand systems, the markup of the firm rises or falls with its market share, and this feature will also hold in our translog system.

In Table 1, we present average Herfindahls at the HS 4-digit level for the U.S. and for the 10 major exporters to the U.S.[4] As one can see from the table, the average U.S. Herfindahl rose slightly over both sub-periods, indicating that increased foreign competition was likely associated with some exit of U.S. firms from the market. If we multiply this average Herfindahl by the share of each country i in the U.S. consumption of good k, s_{it}^{k}, we can compute the typical market share of a firm from that country, $H_{it}^{k}s_{it}^{k}$. We report the weighted average of these per-firm market shares in the last column of Table 1, where the weights are based on the importance of each sector in total U.S. consumption. Table 1 reveals that the share in the U.S. market of a typical U.S. firm fell slightly in the first period and by about 8 percent in the second period. By contrast exporters to the U.S. appear to have gained market share in both periods In other words, those U.S. firms that survived ended up with smaller market shares individually while foreign firms gained market share, which is very much in line with predictions of trade liberalization in the presence of firm heterogeneity as in Melitz and Ottaviano (2008).[5]

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[4] For the U.S., we have adjusted the NAICS 6-digit Herfindahls from the Bureau of Economic Analysis data so that they match the HS 4-digit categories, and detail that procedure in the data Appendix.

[5] Melitz and Ottaviano (2008) describe the equilibrium as follows: "import competition increases competition in the domestic product market, shifting up residual demand price elasticities for all firms at any given demand level. This forces the least productive firms to exit. This effect is very similar to an increase in market size in the closed

One can get a sense of what happened to concentration in other countries by plotting the average export Herfindahl in 2005 against its value in 1992 when we only include sectors for which we could compute a Herfindahl at the HS 4-digit level in both years. The results are shown in Figure 2. The Herfindahl index appears to have risen for most countries in the world indicating that the export market has become more concentrated over time. Nevertheless, the opposite trend seems to be true for many of the most important exporters to the U.S., as listed in Table 1 and labeled in Figure 2. With the exception of Japan, Mexico, and the United Kingdom, all of the remaining top ten exporters to the U.S. saw their export Herfindahls decline over this time period, indicating more exporting firms.

The last row of Table 1 indicates what happened to the average Herfindahl index across markets and countries, as well as the market share of a firm supplying the U.S. market regardless of origin. In order to compute the latter, we multiply each country's average firm's market share by its share in the U.S. market, and sum across countries, obtaining $\sum_i H_{it}^k (s_{it}^k)^2$, which is also averaged across sectors. As one can see from the table, the average firm's market share of an HS 4-digit sector fell by 0.9 percentage points in the first period and 1.9 percentage points in the second. These declines suggest that that market power moderated over both periods.

Obviously, since we cannot measure export Herfindahls in cases where a country does not export, Table 1 and Figure 2 miss one of the most important sources of new competition: the entry of firms into sectors that contained no imports from a particular country previously. Broda and Weinstein (2006) have already extensively documented that this was an important force over

economy: the increased competition induces a downward shift in the distribution of markups across firms. Although only relatively more productive firms survive (with higher markups than the less productive firms who exit), the average markup is reduced. The distribution of prices shifts down due to the combined effect of selection and lower markups. Again, as in the case of larger market size in a closed economy, average firm size and products increase as does product variety. In this model, welfare gains from trade thus come from a combination of productivity gains (via selection), lower markups (pro-competitive effect), and increased product variety."

the period we are examining, so we will not replicate their results except to say that the same forces are at play in our data. Between 1992 and 2005, there was a 54 percent increase in the number of country-HS-10-digit import categories with positive values, which is indicative of substantial foreign entry into new markets. It is the elimination of small U.S. suppliers in the face of the growth of these new foreign suppliers that is the basis of our attempt to quantify the impact foreign entry had on markups, and the number of varieties available for consumption.

This data preview suggests that prior work on the impact of new varieties is likely to suffer from a number of biases. First, as foreign firms have entered the U.S. market there has been exit by U.S. firms, which serves to offset some of the gains of new varieties. Second, while U.S. Herfindahls rose, the Herfindahls of many of our largest suppliers fell. This suggests that there may have been substantial variety growth that is not captured in industry level analyses. Finally, because the market shares of both U.S. firms and the average firm fell over this time period, the rise in foreign entry is likely to have depressed markups overall and therefore lowered prices. Thus, estimates of the gains from new varieties obtained from industry-level data using CES aggregators could either be too large if domestic exit is an important source of variety loss, or too small if foreign firm entry and market power losses are important unmeasured gains. We turn to quantifying these gains and losses in the next section.

3. Translog Function

To introduce the translog function, we will initially simplify our notation above that distinguished countries, firms, and sectors, and instead just let the index i denote products (we will re-introduce countries and firms below). We consider a translog function defined over the universe of products, whose maximum number is denoted by the fixed number $\tilde{N}$. The translog unit-expenditure function is defined by: [6]

$$\ln e = \alpha_0 + \sum_{i=1}^{\tilde{N}} \alpha_i \ln p_i + \tfrac{1}{2} \sum_{i=1}^{\tilde{N}} \sum_{j=1}^{\tilde{N}} \gamma_{ij} \ln p_i \ln p_j \text{, with } \gamma_{ij} = \gamma_{ij} \text{ .} \tag{1}$$

Note that the restriction that $\gamma_{ij} = \gamma_{ij}$ is made without loss of generality. To ensure that the expenditure function is homogenous of degree one, we add the restrictions that:

$$\sum_{i=1}^{\tilde{N}} \alpha_i = 1, \quad \text{and} \quad \sum_{i=1}^{\tilde{N}} \gamma_{ij} = 0 \text{ .} \tag{2}$$

In order to further require that all goods enter "symmetrically" in the γ_{ij} coefficients, we can impose the additional restrictions that:

$$\gamma_{ii} = -\gamma\left(\frac{\tilde{N}-1}{\tilde{N}}\right) < 0, \text{ and } \gamma_{ij} = \frac{\gamma}{\tilde{N}} > 0 \text{ for } i \neq j, \text{ with i, j = 1,}\ldots\text{, } \tilde{N} \text{ .} \tag{3}$$

It is readily confirmed that the restrictions in (3) satisfies the homogeneity conditions (2).

The share of each good in expenditure can be computed by differentiating (1) with respect to $\ln p_i$, obtaining:

$$s_i = \alpha_i + \sum_{j=1}^{\tilde{N}} \gamma_{ij} \ln p_j \text{ .} \tag{4}$$

These shares must be non-negative, of course, but we will allow for a subset of goods to have zero shares because they are not available for purchase. To be precise, suppose that $s_i > 0$ for i=1,…,N, while $s_j = 0$ for j=N+1,…, $\tilde{N}$. Then for the latter goods, we set $s_j = 0$ within the share equations (4), and use these $(\tilde{N} - N)$ equations to solve for the reservation prices $\tilde{p}_j$, j=N+1,…, $\tilde{N}$, in terms of the observed prices p_i, i=1,…,N. Then these reservation prices $\tilde{p}_j$ should appear

[6] The translog direct and indirect utility functions were introduced by Christensen, Jorgenson and Lau (1975), and the expenditure function was proposed by Diewert (1976, p. 122).

within the expenditure function (1) for the unavailable goods j=N+1,…, $\tilde{N}$.

In the presence of unavailable goods, then, the expenditure function becomes rather complex, involving reservation prices. However, if we consider the symmetric case defined by (3), then it turns out that the expenditure function can be simplified considerably, so that the reservation prices no longer appear explicitly. Specifically, Bergin and Feenstra (2009) show that the expenditure function is simplified as:

$$\ln e = a_0 + \sum_{i=1}^{N} a_i \ln p_i + \tfrac{1}{2} \sum_{i=1}^{N} \sum_{j=1}^{N} b_{ij} \ln p_i \ln p_j , \tag{5}$$

where:
$$b_{ii} = -\gamma \frac{(N-1)}{N} < 0, \;\text{ and }\; b_{ij} = \frac{\gamma}{N} > 0 \text{ for } i \neq j \text{ with i, j = 1,…,N,} \tag{6}$$

$$a_i = \alpha_i + \tfrac{1}{N}\left(1 - \sum\nolimits_{j=1}^{N} \alpha_j\right) , \text{ for i = 1,…,N,} \tag{7}$$

$$a_0 = \alpha_0 + \left(\frac{1}{2\gamma}\right)\left\{\sum\nolimits_{i=N+1}^{\tilde{N}} \alpha_i^2 + \left(\frac{1}{N}\right)\left(\sum\nolimits_{i=N+1}^{\tilde{N}} \alpha_i\right)^2\right\}. \tag{8}$$

Notice that the expenditure function in (5) looks like a conventional translog function defined over the available goods i=1,…,N, while the symmetry restrictions in (6) hold analogous to (3), but using the number of available products N rather than $\tilde{N}$. To interpret (7), it implies each of the coefficient α_i is increased by the same amount to ensure that the coefficients a_i sum to unity over i=1,…,N. The final term a_0, appearing in (8), incorporates the coefficients α_i of the unavailable products. If the number of available products N rises, then a_0 falls, indicating a welfare gain from increasing the number of available products. As it is stated, however, (8) does not allow for the direct measurement of welfare gain because it depends on the unknown parameters α_i. We now develop an alternative formula for the welfare gain that depends on the observable expenditures shares on goods, and can therefore be measured.

Let us distinguish two periods t-1 and t, and re-introduce our notation that i denotes countries, while j denotes firms (each selling one good), so the pair (i, j) denotes a unique product variety. We assume that the countries i=M+1,…, $\widetilde{M}$ do not supply in either period, while the countries {1,…,M} are divided into two (overlapping) sets: the M_τ countries $i \in I_\tau$ sell in period τ = t-1,t; with union $I_{t-1} \cup I_t = \{1,...,M\}$ and non-empty intersection $I_{t-1} \cap I_t \neq \emptyset$. We shall let $\bar{I} \subseteq I_{t-1} \cap I_t$ denote a non-empty subset of "common" countries supplying both periods.

Firms in each country provide the set of varieties $j \in J_{it}$, with the number $N_{it} > 0$, so the total number of varieties available each period is $N_t \equiv \sum_{i \in I_t} N_{it}$. If a country supplies in period t but not t-1, then there is obviously an expansion in its set of varieties. But we can also measure an expansion in varieties by examining the Herfindahl indexes of exporting firms for countries supplying *both* periods: a reduction in the Herfindahl indicates greater variety. For our next result, we will need to specify a set of countries $i \in \bar{I}$ for which variety *does not* expand; in practice, we identify these countries by their (relatively) constant Herfindahl indexes. For these countries we assume that there is unchanging sets of variety, $J_{it} \equiv \bar{J}_i$ for $i \in \bar{I}$, with the number $\bar{N}_i > 0$ in each country, so the total number of *unchanging* product varieties is $\bar{N} \equiv \sum_{i \in \bar{I}} \bar{N}_i$.

With this notation, the shares s_{ijt} are now used in place of s_{it} in all our earlier formulas. We can decompose these product shares as $s_{ijt} = s^i_{jt} s_{it}$, where $s_{it} = \sum_{j \in J_{it}} s_{ijt}$ denotes the share of expenditure on all varieties from country i, and $s^i_{jt} \equiv s_{ijt} / s_{it}$ denote the expenditure share on variety j *within* the spending on country i, so that $\sum_{j \in J_{it}} s^i_{jt} = 1$. In practice, we only observe the U.S. import shares s_{it} by country, while we will make inferences about the firm shares s^i_{jt} using the Herfindahl indexes of concentration for each country and product.

Returning to the expenditure function, the Törnqvist price index is exact for the translog function (Diewert, 1976), which means that the ratio of the unit-expenditure functions is measured by:

$$\ln\left(\frac{e_t}{e_{t-1}}\right) = \sum_{i=1}^{M} \sum_{j\in J_i} \tfrac{1}{2}(s_{ijt} + s_{ijt-1})(\ln p_{ijt} - \ln p_{ijt-1}), \tag{9}$$

where $J_i \equiv J_{it} \cup J_{it-1}$ is the set of product varieties sold by country i over both periods. Of course, some of those products may be available in only one period, and likewise, some of the countries i = 1,…,M are selling in only one period. In such cases we again solve for the reservation prices for goods not available, by setting their respective shares equal to zero. Substituting these reservation prices back into (9) and simplifying, we obtain the following expression for the exact price index:

<u>Theorem 1</u>

Then the ratio of translog unit-expenditure functions can be written as:

$$\ln\left(\frac{e_t}{e_{t-1}}\right) = \sum_{i\in \bar{I}} \sum_{j\in \bar{J}_i} \tfrac{1}{2}(\bar{s}_{ijt} + \bar{s}_{ijt-1})(\ln p_{ijt} - \ln p_{ijt-1}) + V, \tag{10}$$

where, the shares $\bar{s}_{ijt-1}$ and $\bar{s}_{ijt}$ are defined as:

$$\bar{s}_{ij\tau} \equiv s_{ij\tau} + \frac{1}{\bar{N}}\left(1 - \sum_{i\in \bar{I}} \sum_{j\in \bar{J}_i} s_{ij\tau}\right), \text{ for } i \in \bar{I} \text{ and } \tau = \text{t-1, t}, \tag{11}$$

and,

$$V \equiv -\left(\frac{1}{2\gamma}\right)\left\{ \sum_{i\notin \bar{I}}(H_{it}s_{it}^2 - H_{it-1}s_{it-1}^2) + \frac{1}{\bar{N}}\left[\left(\sum_{i\notin \bar{I}} s_{it}\right)^2 - \left(\sum_{i\notin \bar{I}} s_{it-1}\right)^2\right]\right\}, \tag{12}$$

where $H_{it} \equiv \sum_{j\in J_{it}} (s_{jt}^i)^2$ denotes the Herfindahl index for firm exports by country i.

Proof: See Appendix

To interpret this result, notice that the constructed shares $\bar{s}_{ij\tau}$ apply to the $\bar{N}$ products that are available in both periods. The constructed shares simply take the observed shares $s_{ij\tau}$ and additively increase each of them by an amount such that $\bar{s}_{ij\tau}$ sum to unity across $\bar{N}$ products. This transformation of shares means that the term $\sum_{i\in\bar{I}}\sum_{j\in\bar{J}_i} \frac{1}{2}(\bar{s}_{ijt}+\bar{s}_{ijt-1})(\ln p_{ijt} - \ln p_{ijt-1})$ appearing in (10) is the Törnqvist price index defined over products available in both periods.[7] The term V defined in (12) is therefore the *extra* impact on the exact price index from having the new and disappearing varieties, and depends on their squared shares, as indicated by the Herfindahl indexes and the country shares s_{it}.

There is one feature of the formula for V that deserves special attention. The first term in curly brackets in (12) is the change in $\sum_{i\notin\bar{I}} H_{it}(s_{it})^2$, summed over those countries not in the set $\bar{I}$. An increase in the Herfindahl index from one of these countries, *ceteris paribus,* would *raise* the variety gain V in absolute value, which is surprising because an increase in the Herfindahl indicates *fewer* exporting firms. The resolution to this puzzle is that the *ceteris paribus* phrase cannot be applied if some exporting firms exit: in that case, there would also be a *fall* in the market share for that country. The formula in (12) must incorporate the change in market share along with the change in the Herfindahl to give an accurate result for welfare.

To illustrate this point with an example, consider the opposite case where there is a rise in the number of suppliers and a fall in the Herfindahl. Specifically, consider a simple example with U.S. consumers purchasing Budweiser and Heineken in period 1, and then having a new domestic variety called American Ale available in period 2.[8] For simplicity, the varieties

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[7] Actually, the $\bar{N}$ products are a subset of those available both periods, since $\bar{I}$ can be a proper subset of the countries i supplying both periods.

[8] American Ale is a new product from the Budweiser company, but we will suppose in our example that this product is being sold by another U.S. firm.

available each period sell in equal shares. The U.S. market share then rises from $s_{us1} = 1/2$ in period 1 to $s_{us2} = 2/3$ in period 2, with Herfindahl indexes $H_{us1} = 1$ in period 1 and $H_{us2} = 1/2$ in period 2 (since then there are two equally sized firms). The change in the U.S. Herfindahl indicates a potential change in variety, so the U.S. is country $i \notin \bar{I}$. In contrast, the Netherlands has unchanged variety (i.e. Heineken), so it is country $i \in \bar{I}$, and so $\bar{N} = 1$. Using this information in (12) we obtain,

$$
\begin{aligned}
V &= -\left(\frac{1}{2\gamma}\right)\left\{\left[\frac{1}{2}\left(\frac{2}{3}\right)^2 - \left(\frac{1}{2}\right)^2\right] + \frac{1}{\bar{N}}\left[\left(\frac{2}{3}\right)^2 - \left(\frac{1}{2}\right)^2\right]\right\} \\
&= -\left(\frac{1}{2\gamma}\right)\left\{\left[-\frac{1}{36}\right] + \frac{1}{\bar{N}}\left[\frac{7}{36}\right]\right\} = -\left(\frac{1}{2\gamma}\right)\frac{1}{6} < 0.
\end{aligned}
$$

The negative value for V lowers the exact price index in (10) and indicates the gain from product variety. Notice that to obtain this negative value, however, we need to incorporate the second term within curly brackets in (12) and above, which is positive; the first bracketed term is negative, reflecting the fall in the U.S. Herfindahl, and on its own would give the wrong sign for the variety gain. So to evaluate V we need to have an accurate value for $\bar{N}$, which in practice we will measure by the sum of the inverse Herfindahls indexes for countries $i \in \bar{I}$, i.e. countries whose Herfindahl indexes do not change by more than some specified tolerance over time.[9]

We conclude with two final observations on V. First, we should not interpret this as the “total” welfare effect of new goods, independently of the Törnqvist index appearing in equation (10). Rather, new goods will also contribute to lower prices for existing goods: this is the pro-competitive effect that we described in the Introduction. Accordingly, we will refer to V as a “partial” welfare effect of new goods; the “total” impact will also have to take into account the pro-competitive effect.

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[9] In our robustness checks we will change the tolerance used to include countries in the “common” set $\bar{I}$ or not.

Second, in order to measure V in (12) we need an estimate of γ. This parameter plays a similar role as the elasticity of substitution in the CES case, in that the welfare gains are reduced as either parameter rises. Obviously, we cannot compare the CES and translog cases without knowledge of these parameters.[10] In both cases, the parameters are estimated from the demand equations. For the translog case, the share equation is obtained by differentiating (5), using (6) and (7), and re-introducing our notation for countries i and firms j, as:

$$s_{ijt} = (\alpha_{ij} + \alpha_t) - \gamma\left(\ln p_{ijt} - \overline{\ln p_t}\right) ,$$

where $\alpha_t = (1 - \sum_{i \in I_t} \sum_{j \in J_{it}} \alpha_{ij})$ is a time-effect which ensures that $\sum_{i \in I_t} \sum_{j \in J_{it}} (\alpha_{ij} + \alpha_t) = 1$, and $\overline{\ln p_t} = \frac{1}{N_t} \sum_{i \in I_t} \sum_{j \in J_{it}} \ln p_{ijt}$ is the average log-price of all available goods in period t.[11]

Using $s_{ijt} = s^i_{jt} s_{it}$ and multiplying the share equation by s^i_{jt}, it becomes:

$$(s^i_{jt})^2 s_{it} = s^i_{jt}(\alpha_{ij} + \alpha_t) - \gamma\left(s^i_{jt} \ln p_{ijt} - s^i_{jt}\overline{\ln p_t}\right) .$$

Summing this equation over $j \in J_{it}$, and noting that $\sum_{j \in J_{it}} s^i_{jt} = 1$, we obtain:

$$H_{it} s_{it} = \alpha_{it} + \alpha_t - \gamma\left(\ln p_{it} - \overline{\ln p_t}\right) , \tag{13}$$

where $\ln p_{it} \equiv \sum_{j \in J_{it}} s^i_{jt} \ln p_{ijt}$ is the (weighted) geometric mean of prices, and $\alpha_{it} \equiv \sum_{j \in J_{it}} s^i_{jt} \alpha_{ij}$ is a (weighted) mean of the taste parameters. This average taste parameter will change as the set of selling firms shifts towards those with higher demand. We therefore model the movement in these tastes parameters as:

[10] Feenstra and Shiells (1997, p. 258) compare the gains from a single new good in the CES and translog cases, by assuming that the new good has the same elasticity of demand in both cases. They show that the "partial" welfare gain from the new good in the translog case is about one-half of the welfare gain in the CES case.

[11] We have included a time subscript on the parameter α_t because it depends on the set of varieties available, which changes over time.

$$\alpha_{it} = \alpha_i + \varepsilon_{it} , \tag{14}$$

where ε_{it} is an error term. Substituting (14) into (13), we obtain the share equations,

$$H_{it}s_{it} = \alpha_i + \alpha_t - \gamma\left(\ln p_{it} - \overline{\ln p_t}\right) + \varepsilon_{it} \ . \tag{15}$$

The parameter γ is obtained by estimating (15), recognizing that the intercept term differs across i and also over time, reflecting changes in the number of available goods. The important properties of these share equation is that the parameter γ *does not* depend on the set of goods available. However, we can expect that the price appearing in (15) are endogenous, as in a conventional supply and demand system. For the CES case, Feenstra (1994) showed how this endogeneity could be overcome without the use of conventional instrument variables, but by exploiting heteroskedasticity in second-moments of the data. We will follow the same procedure in the translog case, as described in section 5. But first, we need to solve for the optimal prices charged by imperfectly competitive firms, in the next section.

4. Optimal Prices and the Pro-Competitive Effect

We will suppose that the available products are produced by single-product firms, acting as Bertrand competitors. The profit maximization problem for firm j in country i is,

$$\max_{p_{ijt}} \ p_{ijt}x_{ij}(p_t, E_t) - C_{ij}[x_{ij}(p_t, E_t)] \ ,$$

where $x_{ij}(p_t, E_t)$ denotes the demand arising from the translog system, with the price vector p_t and expenditure E_t, and $C_{ijt} = C_{ij}[x_{ij}(p_t, E_t)]$ denotes the costs of production. We denote the elasticity of demand by $\eta_{ij}(p_t, E_t) \equiv -\partial \ln x_{ij}(p_t, E_t) / \partial \ln p_{ijt}$ Then the optimal price can be written as the familiar markup over marginal costs:

$$p_{ijt} = C_{ij}'[x_{ij}(p_t, E_t)]\left[\frac{\eta_{ij}(p_t, E_t)}{\eta_{ij}(p_t, E_t) - 1}\right]. \tag{16}$$

The elasticity of demand from the translog system is:

$$\eta_{ijt} = 1 - \left(\frac{\partial \ln s_{ijt}}{\partial \ln p_{ijt}}\right) = 1 + \frac{\gamma(N_t - 1)}{s_{ijt} N_t} \ .$$

It follows that the log-markup appearing in (16) is:

$$\ln\left[\frac{\eta_{ij}(p_t, E_t)}{\eta_{ij}(p_t, E_t) - 1}\right] = \ln\left\{1 + \frac{s_{ijt} N_t}{\gamma(N_t - 1)}\right\}.$$

Substituting these equations into (16), we obtain:

$$\ln p_{ijt} = \ln C'_{ijt} + \ln\left[1 + \frac{s_{ijt} N_t}{\gamma(N_t - 1)}\right], \tag{17}$$

where $C'_{ijt} \equiv C_{ij}'[x_{ij}(p_t, E_t)]$ denotes the time-dependent marginal costs.

We aggregate this equation across firms in each country by multiplying by s^i_{jt} and summing over j:

$$\ln p_{it} = \ln C'_{it} + \sum_{j \in J_{it}} s^i_{jt} \ln\left[1 + \frac{(s^i_{jt} s_{it}) N_t}{\gamma(N_t - 1)}\right], \tag{18}$$

where $\ln p_{it}$ is again the geometric mean of prices, and $\ln C'_{it} \equiv \sum_{j \in J_{it}} s^i_{jt} \ln C'_{ijt}$ is the geometric mean of marginal costs in country i. In order to evaluate this expression, we need to bring the summation (like an expectation) within the log expression, which means that we are ignoring Jensen's inequality; we argue below that this is a second-order approximation.[12] In that case, we obtain the final form of our pricing equation:

[12] Note that we need to make this approximation due to missing data at the firm level. For the United States, for example, we have Herfindahl indexes at the sector level but not the underlying firm shares.

$$\ln p_{it} = \ln C'_{it} + \ln\left[1 + \frac{\sum_{j\in J_{it}} s^i_{jt}(s^i_{jt}s_{it})N_t}{\gamma(N_t-1)}\right] \approx \ln C'_{it} + \ln\left[1 + \frac{H_{it}s_{it}N_t}{\gamma(N_t-1)}\right]. \qquad (19)$$

The pro-competitive effect is obtained by substituting the pricing equation (19) into (10). The resulting expression involves both share-weighted and unweighted geometric means of the firm prices, because the shares $\bar{s}_{ij\tau}$ in (11) are additive transformations of the shares $s_{ij\tau}$. In practice we will not be able to distinguish weighted and unweighted firm prices, and simply use import unit-values for either. So to eliminate this distinction in the theory, we strengthen our earlier assumption that countries supplying in both periods have unchanging sets of variety, $J_{it} \equiv \bar{J}_i$ for $i \in \bar{I}$. Specifically, we now assume that if there is no entry or exit of firms in a country, then the firm shares are *equal and unchanging* within that country:

$$s^i_{j\tau} = \frac{1}{\bar{N}_i}, \text{ for } i \in \bar{I},\ \tau = \text{t-1,t}. \qquad (20)$$

Notice that the *country shares* s_{it} still change for countries selling in both periods, so that (20) specifies that firms within these countries $i \in \bar{I}$ do not change size relative to their country sales. In that case, the pro-competitive effect is written as follows:

<u>Theorem 2</u>

For $i \in \bar{I}$, the pricing equation (19) is a second-order approximation to (18) around the point where $s_{it}N_t/\gamma(N_t-1)=0$ and (20) holds. Then using (19) and (20), the pro-competitive effect P is:

$$\ln\left(\frac{e_t}{e_{t-1}}\right) = \sum_{i\in\bar{I}} \tfrac{1}{2}(\bar{s}_{it} + \bar{s}_{it-1})(\ln C'_{it} - \ln C'_{it-1}) + V + P,$$

with the shares $\bar{s}_{i\tau} \equiv s_{i\tau} + \frac{\bar{N}_i}{\bar{N}}\left(1 - \sum_{j\in\bar{I}} s_{j\tau}\right)$, for $i \in \bar{I}$ and τ = t-1, t, and,

$$P \equiv \sum_{i\in \bar{I}} \tfrac{1}{2}(\bar{s}_{it} + \bar{s}_{it-1})\left\{ \ln\left[1+\frac{H_{it}s_{it}N_t}{\gamma(N_t - 1)}\right] - \ln\left[1+\frac{H_{it-1}s_{it-1}N_{t-1}}{\gamma(N_{t-1}-1)}\right]\right\}. \tag{21}$$

Using $\ln(1+x) \approx x$, the pro-competitive effect is approximated as:

$$P \approx V + \left(\frac{1}{2\gamma}\right)\sum_{i=1}^{M}(H_{it}s_{it}^2 - H_{it-1}s_{it-1}^2) + \frac{1}{2\gamma}\sum_{i\in \bar{I}}(\bar{s}_{it} + \bar{s}_{it-1})\left[\frac{H_{it}s_{it}}{(N_t - 1)} - \frac{H_{it-1}s_{it-1}}{(N_{t-1}-1)}\right]. \tag{22}$$

Proof: See Appendix

Equation (21) is the final form for the pro-competitive impact that we will evaluate, while (22) provides use with some intuition on this term: the pro-competitive effect lowers the exact price index by more that the partial variety effect whenever the additional terms on the right of (22) are negative. Focusing on the second term on the right, we can see that the pro-competitive effect lowers the price index by more than the partial variety effect provided that $\sum_{i=1}^{M} H_{it}s_{it}^2$ is falling over time. It is useful to give a more precise interpretation to that term. Recalling that the Herfindahl indexes are $H_{it} \equiv \sum_{j\in J_{it}} (s_{jt}^i)^2$, we see that:

$$\sum_{i=1}^{M} H_{it}s_{it}^2 = \sum_{i=1}^{M}\sum_{j\in J_{it}} (s_{jt}^i)^2 s_{it}^2 = \sum_{i=1}^{M}\sum_{j\in J_{it}} s_{ijt}^2 \equiv H_t^k \,. \tag{23}$$

In words, the sum of the Herfindahl firm indexes weighted by the squared country shares, on the left of (23), is exactly the right way to aggregate these indexes to obtain an *overall Herfindahl for the good k in question,* on the right of (23). This summary statistic was shown in the last row and column of each panel in Table 1, when averaged across sectors. We therefore see that a falling *overall* Herfindahl contributes to lowering prices through the pro-competitive impact, as we suggested in our data preview in section 2.

5. Estimation and Results

We turn now to estimation of the translog parameter γ. We will specify that the weighted average of marginal costs from each exporting country take on the iso-elastic form:

$$\ln C'_{it} = \omega_{i0} + \omega \ln\left(\frac{s_{it}E_t}{p_{it}}\right) + \delta_{it} \,,$$

where the term $(s_{it}E_t / p_{it})$ reflects the total quantity exported from country i, and δ_{it} is an error term. Substituting into (19), we obtain a modified pricing equation:

$$(1+\omega)\ln p_{it} = \omega_{i0} + \omega \ln s_{it} + \omega \ln E_t + \ln\left[1 + \frac{H_{it}s_{it}N_t}{\gamma(N_t - 1)}\right] + \delta_{it} \,. \qquad (24)$$

We see that the translog parameter γ appears in both the share equation (15) and the pricing equation (24): larger γ means that the goods are stronger substitutes and the markups are correspondingly smaller. It is also evident that the shares and prices are endogenously determined: shocks to either supply δ_{it} or demand ε_{it} will both be correlated with shares s_{it} and prices p_{it}. To control for this endogeneity will we estimate these equations simultaneously using a similar methodology to that proposed in the CES case by Feenstra (1994) and extended by Broda and Weinstein (2006).

The first step in our estimation is to difference (15) and (24) with respect to country k and with respect to time, thereby eliminating the terms $\alpha_i + \alpha_t$ and the overall average prices $\overline{\ln p_t}$ appearing in the share equations, and eliminating total expenditure $\ln E_t$. We also divide the share equation by γ and the pricing equation by $(1+\omega)$, and then express each equation in terms of its error term:

$$\frac{(\Delta\varepsilon_{it} - \Delta\varepsilon_{kt})}{\gamma} = \frac{[\Delta(H_{it}s_{it}) - \Delta(H_{kt}s_{kt})]}{\gamma} + (\Delta \ln p_{it} - \Delta \ln p_{kt}) \,,$$

$$\frac{(\Delta\delta_{it}-\Delta\delta_{kt})}{(1+\omega)}=(\Delta\ln p_{it}-\Delta\ln p_{kt})-\frac{\omega(\Delta\ln s_{it}-\Delta\ln s_{kt})}{(1+\omega)}$$
$$-\frac{1}{(1+\omega)}\left\{\Delta\ln\left[1+\frac{H_{it}s_{it}N_t}{\gamma(N_t-1)}\right]-\Delta\ln\left[1+\frac{H_{kt}s_{kt}N_t}{\gamma(N_t-1)}\right]\right\}.$$

We multiply these two equations together, and average the resulting equation over time, to obtain the estimating equation:

$$\overline{Y}_i=\frac{\omega}{(1+\omega)}\overline{X}_{1i}+\frac{\omega}{\gamma(1+\omega)}\overline{X}_{2i}-\left(\frac{1}{\gamma}\right)\overline{X}_{3i}+\frac{1}{(1+\omega)}\overline{Z}_{1i}(\gamma)+\frac{1}{\gamma(1+\omega)}\overline{Z}_{2i}(\gamma)+\overline{u}_i, \quad (25)$$

where the over-bar indicates that we are averaging that variable over time, and:

$$Y_{it}\equiv(\Delta\ln p_{it}-\Delta\ln p_{kt})^2,$$

$$X_{1it}\equiv(\Delta\ln s_{it}-\Delta\ln s_{kt})(\Delta\ln p_{it}-\Delta\ln p_{kt}),$$

$$X_{2it}\equiv(\Delta\ln s_{it}-\Delta\ln s_{kt})[\Delta(H_{it}s_{it})-\Delta(H_{kt}s_{kt})],$$

$$X_{3it}\equiv(\Delta\ln p_{it}-\Delta\ln p_{kt})[\Delta(H_{it}s_{it})-\Delta(H_{kt}s_{kt})],$$

$$Z_{1it}(\gamma)\equiv\left\{\Delta\ln\left[1+\frac{H_{it}s_{it}N_t}{\gamma(N_t-1)}\right]-\Delta\ln\left[1+\frac{H_{it}s_{kt}N_t}{\gamma(N_t-1)}\right]\right\}(\Delta\ln p_{it}-\Delta\ln p_{kt}),$$

$$Z_{2it}(\gamma)\equiv\left\{\Delta\ln\left[1+\frac{H_{it}s_{it}N_t}{\gamma(N_t-1)}\right]-\Delta\ln\left[1+\frac{H_{it}s_{kt}N_t}{\gamma(N_t-1)}\right]\right\}(\Delta H_{it}s_{it}-\Delta H_{kt}s_{kt}).$$

and, $$u_{it}\equiv\frac{(\Delta\varepsilon_{it}-\Delta\varepsilon_{kt})(\Delta\delta_{it}-\Delta\delta_{kt})}{\gamma(1+\omega)}.$$

We shall assume that the error terms in demand and the pricing equation are uncorrelated, which means that the error term in (25) becomes small, $\overline{u}_i\to 0$ in probability limit as $T\to\infty$. That error term is therefore uncorrelated with any of the right-hand side variables as $T\to\infty$, and we can exploit those moment conditions by simply running OLS on (25). Feenstra (1994) shows that procedure will give us consistent estimates of γ and ω in a slightly simpler system, provided that the right-hand side variables in (25) are not perfectly collinear as $T \rightarrow \infty$. As in the CES case of Feenstra (1994), that condition will be assured if there is some heteroskedasticity in the error terms across countries i, so that the right-hand side variables in (25) are not perfectly collinear. More efficient estimates can be obtained by running weighted least squares on (25).

Before proceeding with the estimation, we had to solve a number of data problems. First, while in principle we could have estimated γ at the 10-digit level, in practice this is not possible because we do not have enough 10-digit varieties in most sectors. In order to make sure that we had enough data to obtain precise estimates, we decided to assume that the γ's at the 10-digit level within an HS-4-digit sector were the same. This assumption meant that we typically had 99 varieties when we estimated a γ for an HS-4 sector.

A second complication arises because we have U.S. shipments data at the NAICS-6 digit level but we need to compute shares at the HS-10 digit level. Thus, we had to allocate NAICS-6 production data to each HS-10 sector. In order to do this, we assumed that the share of U.S. production in each HS-10 was the same as that of the U.S. in the NAICS-6 digit sector that contains it, as discussed in the data Appendix.

A third complication arises because we use unit-values of import prices from each source country rather than the geometric mean, which introduces measurement error, especially for import flows that are very small. Broda and Weinstein (2006) propose a weighting scheme based on the quantity of imports at the HS-10 level. Unfortunately, we could not implement precisely that scheme because the U.S. quantity indexes were defined at the NAICS-6 digit level and not at the HS-10 digit level. We therefore decided to implement the Broda and Weinstein weighting scheme using value of shipments instead of quantity of shipments, since shipment values are likely to be highly correlated with shipment quantities across countries.

Finally, as in Broda and Weinstein (2006), we also faced the problem that only 86 percent of our estimates of γ had the right sign if we estimate them without constraints. If γ is less than zero, then this implies that demand is inelastic and the welfare gains associated with new and disappearing varieties are infinite. Since we wanted to rule this out and because the formula for V is very sensitive to small values of γ, we decided to place a constraint on γ limiting it to have a smallest value of 0.05. In order to do this, we used a grid search procedure over γ and ω to minimize the sum squared errors in equation (25). In this procedure we set an initial γ of 0.05 and increased it by 5 percent over the range [.05, 110].[13] Similarly, we set an initial ω of -5 and increased it by 0.1 over the range [-5, 15].

Plots of the Data

Equations (12) and (21) or (22) are the key equations for understanding how new varieties affect consumers through increased choice and lower markups. Before we present the final results, it is worth going through a decomposition of the components so that we can understand the forces at play.

We begin with the partial variety effect, V, in (12). Its coefficient $1/2\gamma$ captures the fact that consumers care more about goods that are less elastically demanded (i.e. have low γ's) than goods that have close substitutes. The term in curly brackets in (12) can be understood by breaking it up into its components. First, $H_{i\tau}s_{i\tau}$ is the typical firm's market share. In order to compute the aggregate impact of variety creation and destruction, we need to aggregate these, but the aggregation process places more weight on goods that have higher market shares than

[13] In order to speed up the grid searches, in most specifications we increased the interval by 5 percent until 7.8 and then jumped to 109.9. We did this because we almost never found gammas between 7.8 and 109.9. Moreover, making this change did not qualitatively affect the results because all high gammas imply very small markups and variety effects.

those with lower shares. As a result, we aggregate these across varieties by weighting them by $s_{i\tau}$ and create $\sum_{i \notin \bar{I}} H_{i\tau} s_{i\tau}^2$. Essentially, the partial variety effect will have a negative effect on the price level if the market share of new entrants is, on average, larger than that of firms that exit. Likewise, the second term in curly brackets measures the country share of new suppliers versus disappearing supplies. Thus, equation (12) indicates that the partial variety effect will be driven by how important new varieties are in demand.

Before we turn to the estimation, we can obtain some intuition for the expected results by plotting the distribution of $2\gamma V$, which corresponds to the negative of the term in curly brackets in (12). Since $\gamma > 0$, the sign of the variety gain, V, will be the same as the sign of $2\gamma V$ but requires no estimation. Since we simply observe Herfindahls and not firm-level data, we decided to define a new variety as the appearance or disappearance of an HS-10 digit export from a country or whenever the Herfindahl in 2005 relative to that in 1992 fell within the range of [1/1.3, 1.3]. We will explore the robustness of our results to this criterion later, but this seems like a reasonable starting point. Thus, the countries $i \in \bar{I}$ in each sector are those that export the U.S. in both years and have the Herfindahl ratio in that range. Figure 3a plots the distribution of $2\gamma V$. The distribution is fairly symmetric although there appears to be a slightly negative mass.[14] Indeed, both the median and the mean are negative (-0.03 and -1.1 percent).

We can use a similar technique to understand the distribution of the pro-competitive effect, P. If we multiply both sides of equation (22) by $2\gamma > 0$, we can write $2\gamma P$ as a function of $2\gamma V$ and two terms that are composed of Herfindahl indexes. The second of these terms is a decreasing function of the number of firms in the sector. If we assume the number of firms is

[14] There are a few larger positive and negative outliers that we do not show in any version of Figure 3 because they would compress the distribution too much. All of our results are robust to dropping the top and bottom 1 percent of the V distribution.

large, then this term will be close to zero and we can ignore it for now (even though we will include it when we compute P in the next section). This simplification enables us to now write 2γP as purely a function of the raw data. We plot this distribution in Figure 3b. As one can see from the histogram, the distribution of 2γP is much more sharply shifted to the left. The median and mean are – 0.05 and – 1.2 percent respectively, suggesting fairly substantial pro-competitive effects (as long as γ is not too large).

Finally, Figure 3c plots the distribution of $\sum_{i=1}^{M}(H_{it}s_{it}^{2} - H_{it-1}s_{it-1}^{2})$ which corresponds to 2γ(P – V), once again assuming for the moment the last term in equation (22) is approximately zero. This value tells us how changes in market Herfindahls alone affected markups. Again the mass of this distribution is greater to the left of the zero indicating that firm market shares fell on average during this period. This suggests that the decline in the typical firm's market share put downward pressure on prices.

Thus, even before we turn to the estimation, the data suggests that consumers of merchandise were likely to have benefited from increased variety, as indicated by the sign of 2γV, and a decline in markups, as indicated by the sign of 2γP. In order to understand the impact of these changes on welfare, however, we need to estimate γ for each sector and aggregate.

Estimation Results

Because we ultimately estimated over one thousand γ's, it is not possible to display all of them here. We display the sample statistics for γ in Table 2. The median γ was 0.19 and the average was 12. The large average γ is driven by the fact that their distribution is not symmetric and γ can take on very large values. It is difficult to have strong priors for what a reasonable value of γ should be. One way possible benchmark is the implied markup. We can compute the markup for each industry by using equation (19). Based on this calculation the median estimated markup in our data is 0.30 (i.e. a 30% markup over marginal costs) in 2005. By comparison, Domowitz, Hubbard, and Petersen (1988) estimate markups across U.S. manufacturing and obtain an average markup of 0.37, which is a bit higher than ours but not dramatically different especially given the large differences in data and estimating procedures.

The markups in each sector depend on the value of the firm's market share as well. We can get some sense of the reasonableness of our estimates by looking at the most important sectors in U.S. absorption. In Table 3, we report the share of U.S. absorption from the ten largest sectors (with names not beginning with "other"), where we define the share to be the average share of absorption in 1992 and 2005. In the first column we report our estimate of γ. Based on this measure, we find the three sectors where the products are most heterogeneous and firms are likely to have the most market power are "Aircraft and Spacecraft," "Televisions, Video Cameras, and Receivers," and "Private Motor Vehicles." In contrast, the most homogeneous sectors where firms are likely to have the least market power are "Crude Petroleum," "Natural Gas," and "Cigarettes and Cigars." This pattern seems broadly sensible.

We now are ready to present aggregate estimates of P and V for all merchandise consumed in the U.S. In order to do this, we aggregated P and V computed at the HS-4 level using the formula:

$$\hat{P} = \sum_k \tfrac{1}{2}(s_{kt} + s_{kt-1})P_k \text{ and } \hat{V} = \sum_k \tfrac{1}{2}(s_{kt} + s_{kt-1})V_k \tag{26}$$

where we reintroduce the sector subscript k, and hence P_k and V_k are the values for P and V computed at the HS-4 level and $s_{k\tau}$ is the share of that sector in U.S. absorption. Our baseline estimate for P and V are -0.017 and -.037, which means that the welfare gain due to the decline in markups is 1.7 percent and the partial gain from varieties is about 3.7 percent. These numbers are recorded in the first row of Table 4. Thus, the combined impact is to lower the U.S. merchandise price index by 5.4 percent between 1992 and 2005. Given that U.S. merchandise demand constituted 18.5 percent of GDP in 2002, this corresponds to a 1.0 percent gain for U.S. consumers.[15] Of this gain, 0.31 percentage points comes from lower markups and the remaining 0.69 percentage points comes from the partial variety effect.

We can obtain some intuition for these numbers by returning to the results we presented in the discussion of Figures 4a and 4b. There, we found that the mean value of $2\gamma V$ was -1.1 percent and the mean value of $2\gamma P$ was -1.2 percent. If we simply apply our median estimates of γ to these numbers, we would obtain a partial variety impact on prices of -2.9 percent and partial markup effect of -3.2 percent and therefore an aggregate impact of -6.1 percent. This is somewhat larger than the -5.4 percent we estimate and suggests that full distribution of γ's serves to lower our point estimate of the welfare gains due to new varieties and lower markups, but that our results are not being driven by an outlier value of γ.

Obviously one concern is the precision of our estimates. Because of the nonlinearity and our grid search algorithm, computation of the confidence intervals for P and V is not straightforward. We decided to compute these by bootstrapping each of the 1000 γ's and ω's and then using these bootstrapped parameter values to compute the distribution of P and V. This is enormously computationally intensive, but ultimately we were able to compute P and V 100 times.[16] In our baseline case, we found that the 10-90 percent confidence interval for P was [-0.020, -.015] while the same interval for V was [-.056, -0.18]. This indicates that our point estimates for the markup and variety effects are estimated with reasonable precision.

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[15] We define merchandise demand as U.S. GDP in agriculture, mining, and manufacturing less exports plus imports in those sectors.

[16] It took 10 days on an 8-processor SPARCstation.

Robustness Tests

One surprising feature of our welfare estimates so far is the relative ranking of P and V. Following Theorem 2 we argued that if $\sum_{i=1}^{M} H_{it} s_{it}^2$ is falling, indicating that sector k is becoming more competitive, then we should have $P_k < V_k < 0$, so the pro-competitive effect is more important than the partial variety effect in lowering prices. Averaging over sectors we found that $\sum_{i=1}^{M} H_{it} s_{it}^2$ was indeed falling, as shown in Table 1 (entry in last row and column of each panel) and in Figure 3c (which has a negative mean). But when we apply the estimates of γ_k in each sector to compute P_k and V_k, and then average across sectors, we find that $\hat{V} < \hat{P} < 0$, so the variety effect exceeds the pro-competitive effect.

To explain this reversal of the ranking, we looked at which U.S. sectors were causing most of the variety gains. We find that of the 2.9 percentage point drop in merchandise prices due to new varieties, 2.4 percentage points was due to a single sector: new automobile and truck varieties. Between 1992 and 2005, there was enormous entry into this sector as Japanese car makers set up new plants (see Blonigen and Soderbery (2009)). This entry had two important impacts. First, the U.S. Herfindahl index declined sharply from 0.35 to 0.21, reflecting the large increase in the number of makers operating in the U.S. Second, the transplant of Japanese car makers to the U.S. was associated with a very large increase in U.S. automobile production: real output of autos made in the U.S. grew by 41 percent between 1992 and 1998, which contributed to a substantial increase in the share of U.S. consumption made domestically. That increase in the share resulted in a very large welfare gain, or drop in V, from equation (12).

There are reasons to believe, however, that our welfare formula cannot accurately deal with the transplant of Japanese varieties to the United States: we have ignored multi-product firms, for example, and in the same way have assumed that the γ estimate for autos applies equally well to products *across* firms as to products *within* firms. That assumption clearly contradicts the theoretical literature on multi-product firms, which makes a strong distinction between consumer substitution of products within and between firms (see Allanson and Montagna, 2005, and Bernard, *et al* 2006a,b). For this reason we also computed the aggregate values for $\hat{P}$ and $\hat{V}$ while ignoring the passenger vehicle sector. This gives the result shown in the second row of Table 4, where both the pro-competitive and partial variety gains are 0.13, or welfare gains of 1.3 percentage points each. The sum of these is only one-half as big as our benchmark estimates, and now the pro-competitive and partial variety effects are of roughly equal magnitude.

In Table 4 we present some additional robustness tests of estimated impacts. The next robustness check consists of varying the sensitivity of the estimates to the cutoff Herfindahl we use to determine whether a country is in the set $\bar{I}$ or not, i.e. whether it is a "common' country in both time periods with unchanged exporting firms. In our baseline case we examined fluctuations in the Herfindahl of [1/1.3, 1.3], but we also examined fluctuations of as tight as [1/1.1, 1.1] and as loose as [1/1.5, 1.5]. These are reported in the next four rows of Table 4. Theoretically, our results in Theorems 1 and 2 hold for any nonempty set $\bar{I} \subseteq I_t \cap I_{t-1}$, and for this reason we might expect our aggregate estimates to be invariant to this $\bar{I}$ cutoff. In practice, the sensitivity arises for a number of reasons. First, as we tighten the Herfindahl criterion we lose some sectors because we no longer have any "common" countries, and so Theorem 1 cannot be applied in those sectors. For example, there are 21 more sectors when we use a criterion of [1/1.3, 1.3] than when we use [1/1.1, 1.1]. Secondly, our estimation relies on the assumption that the number of firms equals the inverse of the Herfindahl, which is not exact.

The fact that V tends to rise and then fall in Table 4 as we change the cutoff can be explained by looking at the second term in curly brackets in (12). If we have a very tight cutoff for "common" countries, then a large share of the trade flows will be new or disappearing and the corresponding shares of new and disappearing goods in the second term will *both* approach one. Thus V will tend to be small because the difference between two squared share terms will approach zero. If the cutoff is very loose, however, then this will mean that the number of common firms, $\bar{N}$, will be large and V will also tend to be small. Thus, the partial variety effect is dependent on the cutoff we choose. Still, it is remarkable how stable our estimates are to variation whether we count countries as entering or exiting particular sectors.

We also wanted to ensure that our efforts to concord the SIC sectors with the NAICS sectors was not driving the results. To check this, we split the sample into two periods 1992-1997 and 1998-2005 and reran the estimation for each period. We then summed the markup and variety effects estimated over each period and report them in the second-last row of Table 4. As one can see from the table, whether we use the merged data or work with different subsamples does not have a large impact on our results. We still obtain an aggregate impact on merchandise prices of 5 percent, of which two thirds is driven by variety gains (including passenger vehicles).

Comparison with CES Case

Our baseline estimate of the impact of new goods and changing markups on prices is 5.4 percent, although depending on the cutoff for common goods this estimate can be as low as 3.9 percent (or 2.6 percent without autos), as shown in the first various rows of Table 4. The magnitudes of these numbers are perhaps easiest to understand relative to Broda and Weinstein's (2006) estimates for the period 1990 to 2001. Those authors used a CES aggregator and obtained a gain to consumers of 0.8 percent over the 1990-2001 period. That is slightly larger than the

0.69 benchmark percent estimate of pure variety gain in this paper, and slightly below our aggregate estimate of 1.0 percent. But the two estimates are not directly comparable for three reasons: first, Broda and Weinstein used both a different functional form (CES); second, they assumed that there was no firm entry or exit in sectors in which a country exported in the beginning and end of the sample; and third, they estimated the gain over 11 years instead of 13. If we multiply our estimates by 11/13, we find that implied aggregate gain over an 11-year period in the translog case is also 0.8. This suggests that both functional forms yield surprisingly similar aggregate welfare gains.

Nevertheless there are some important differences. In particular, while the CES aggregator ascribes all of the welfare gain to new varieties, the partial impact of new varieties in the translog case is 69 percent of the total gains in our benchmark translog estimates, or one-third smaller than in the CES case. We can obtain some sense of how important the pure functional form assumptions are by setting the Herfindahls of all countries equal to their 1992 values and recalculating the variety gain. In this case, we are assuming, as in Feenstra (1994) and Broda and Weinstein (2006), that the only source of new varieties is the entry and exit of *exporting countries* in each product market. Eliminating the impact of firm entry and exit within sectors gives us a variety impact on the price level of 0.4 percent – only an eighth as large as before – as shown in the last row of Table 4. While this causes measured variety gains to be much smaller, it also causes measured drops in markups to be much larger because there is no exit in response to foreign entry. When we fix Herfindahls to their 1992 level, the aggregate drop in prices is 3.2 percent – almost double what is was before. In other words, the translog functional form ascribes a smaller role for variety than the CES, but we obtain comparable results in this paper to that of Broda and Weinstein (2006) because variety growth also has important impacts on markups in a translog setup that are not permitted in the CES framework.

6. Conclusions

Krugman (1979) demonstrated the reduction in markups that accompanies trade liberalization under monopolistic competition. That reduction in markups is not just a consumer gain, but is also a social gain: the reduction in markups in a zero-profit equilibrium indicates that the wedge between firm’s marginal and average costs is reduced, so that output is expanding and there are greater economies of scale. So the competition between firms from different countries is an important channel by which international trade leads to social gains.

Despite this insight, such a channel has received only limited attention in the empirical trade literature. We have argued that the reason for this gap in the literature is the common assumption of CES preferences, which leads to constant markups. So instead we must look to alternative preferences, of which the quadratic preferences in Melitz and Ottaviano (2008) are a leading example. On empirical grounds we have adopted instead translog preferences, simplified to impose symmetry in substitution across products. We have derived quite general formulas for the welfare gains from new products with these preferences, and also the pro-competitive effect of new entry on reducing markups. These formulas allow for multiple countries with firms that are heterogeneous in their marginal costs, and nearly any pattern of exit and entry, subject to the identifying assumption that some countries have unchanged sets of firms over time.

The translog preferences lead to log-linear demand and pricing equations, which we estimate jointly. In this respect we are following the general approach of the industrial organization literature (Bresnahan, 1989; Berry, 1994): markups are not observed directly because marginal costs are not observed, so we rely on estimates of the elasticity of demand to identify the markups. But unlike the industrial organization literature, we are not interested here in a single market, but rather, in estimating the impact of globalization on markups for an entire economy – the United States. To address the simultaneity of supply and demand across so many markets, we rely on the "identification through heteroskedasticity" approach used by Feenstra (1994) and Broda and Weinstein (2006), and extended here from the CES to the translog setting.

The tremendous amount of entry of foreign countries into U.S. markets, as well as more exporters within those countries, drives our measure of the variety gains. This entry has been offset to some degree by the exit of firms from the United States, leading to a rise in those Herfindahl indexes. Nevertheless, we find that the exit from the U.S. market has been less than the new entry, in the sense that the rise in U.S. Herfindahls is less that the fall the overall U.S. share, so that the *per-firm* share of surviving U.S.. firms fell in many sectors. That feature of the data drives our estimates of the fall in markups, which is the pro-competitive effect of globalization.

In our benchmark results, we find that *total* welfare gain from globalization for the U.S. in the translog case is of the same magnitude as that found by Broda and Weinstein (2006) in the CES case, but that the composition of this gain is different. In theory, we could expect the pro-competitive effect to be larger than the welfare gain from new varieties, but in our benchmark estimates the opposite ranking occurs: the pro-competitive effect was about one-third of the total gain and the variety effect was two-thirds. But that result is sensitive to one very large sector, passenger motor vehicles, without which the two sources of gain are about equal in size. So we conclude that while translog preferences give variety gains at least one-third lower than in the CES case, the additional pro-competitive effect can plausibly lead to similar overall gains from globalization under the two functional forms.

Appendix A: Proofs of Theorems

Proof of Theorem 1:

For convenience we denote the firm-country pairs (i,j) instead by just the product index i, where products i=1,…,N are available in period t-1 or t. These are divided into two (overlapping) sets: the products $i \in I_\tau$ sell in period τ = t-1,t; with their union $I_{t-1} \cup I_t = \{1,...,N\}$ and non-empty intersection $I_{t-1} \cap I_t \neq \emptyset$. We shall let $\bar{I} \subseteq I_{t-1} \cap I_t \neq \emptyset$ denote any non-empty subset of their intersection, and without loss of generality we order the goods so that the first N_1 goods denoted i=1,…, N_1 are in $\bar{I}$, and therefore available both periods (N_1 equals $\bar{N}$ as used in the text); while the next N_2 goods denoted i= N_1+1,…,N are available in either one or both periods, but are not in $\bar{I}$. These two categories exhaust the N goods, N= N_1+N_2. The expenditure function is as shown in equations (5) – (8), and Törnqvist price index is,

$$\ln\left(\frac{e_t}{e_{t-1}}\right) = \sum_{i=1}^{N} \tfrac{1}{2}(s_{it} + s_{it-1})(\ln p_{it} - \ln p_{it-1}).$$

Let B denote the NxN matrix B = $-\gamma I_N + (\gamma / N)L_{NxN}$, where I_N is the NxN identity matrix and L_N is an NxN matrix with all elements equal to unity. We partition the B matrix into the same two mutually exclusive groups, and likewise for the vector a:

$$a = \begin{bmatrix} a^1 \\ a^2 \end{bmatrix}, \text{ and } \quad B = \begin{bmatrix} B^{11} & B^{12} \\ B^{21} & B^{22} \end{bmatrix}.$$

The diagonal elements in the matrix B are $B^{kk} = -(\gamma / N)[N\ I_{N_k} - L_{N_k x N_k}]$, and the off-diagonal elements are $B^{12} = B^{21}{}' = (\gamma / N)[L_{N_1 x N_2}]$. Similarly, we partition the share vectors $s_\tau^1 = (s_{1\tau},...,s_{N_1,\tau})'$ and $s_\tau^2 = (s_{N_1+1,\tau},,...,s_{N\tau})'$, and likewise for the price vectors p_τ^1 and p_τ^2, τ=t-1,t. If $\bar{I} = I_{t-1} \cap I_t$, then all the goods i= N_1+1,…,N are new or disappearing , with either

$s^2_{it-1} = 0$ or $s^2_{it} = 0$. More generally, with $\bar{I} \subset I_{t-1} \cap I_t$ then *some* of the goods i= N_1+1,…,N are new or disappearing, with zero share. So we use the notation $\tilde{p}^2_\tau$ to denote the reservation prices for those goods with zero share in period τ = t-1,t, but the same vector uses *actual* prices for those goods with positive shares.

Then the share equations in periods t-1 and t for the goods i= N_1+1,…,N are:

$$s^2_{t-1} = a^2 + B^{21} \ln p^1_{t-1} + B^{22} \ln \tilde{p}^2_{t-1},$$

$$s^2_t = a^2 + B^{21} \ln p^1_t + B^{22} \ln \tilde{p}^2_t,$$

where some of these shares can be zero. From these equations we solve for the reservation prices for new and disappearing goods (and actual prices for the goods with positive shares):

$$B^{22} \ln \tilde{p}^2_{t-1} = (s^2_{t-1} - a^2 - B^{21} \ln p^1_{t-1}),$$

and,

$$B^{22} \ln \tilde{p}^2_t = (s^2_t - a^2 - B^{21} \ln p^1_t).$$

It follows that,

$$(\ln \tilde{p}^2_t - \ln \tilde{p}^2_{t-1}) = \left[B^{22}\right]^{-1}\left[(s^2_t - s^2_{t-1}) - B^{21}(\ln p^1_t - \ln p^1_{t-1})\right]. \tag{A1}$$

Substituting (A1) into the Törnqvist price index, we obtain:

$$\begin{aligned}\ln\left(\frac{e_t}{e_{t-1}}\right) &= \tfrac{1}{2}(s^1_t + s^1_{t-1})'(\ln p^1_t - \ln p^1_{t-1}) - \tfrac{1}{2}(s^2_t + s^2_{t-1})'\left[B^{22}\right]^{-1} B^{21}(\ln p^1_t - \ln p^1_{t-1}) \\ &+ \tfrac{1}{2}(s^2_t + s^2_{t-1})'\left[B^{22}\right]^{-1}(s^2_t - s^2_{t-1})\end{aligned} \tag{A2}$$

From the definition of the partitioned matrix B, we have that:

$$B^{22} = -\left(\frac{\gamma}{N}\right)[N\ I_{N_2} - L_{N_2 \times N_2}], \tag{A3}$$

where $[N\ I_{N_2} - L_{N_2 \times N_2}]$ has an eigenvector $L_{N_2 x1}$ with the associated eigenvalue of N_1, so its inverse matrix has the reciprocal eigenvalue. Then by definition of $B^{21} = (\gamma / N)[L_{N_2 x N_1}]$ we can simplify the second term on the right of (A2) as:

$$-\tfrac{1}{2}(s_t^2+\tilde{s}_{t-1}^2)'\left[B^{22}\right]^{-1}B^{21}(\ln p_t^1-\ln p_{t-1}^1)$$

$$=\tfrac{1}{2}(s_t^2+s_{t-1}^2)'\left(\frac{N}{\gamma}\right)[N\,I_{N_2}-L_{N_2xN_2}]^{-1}B^{21}(\ln p_t^1-\ln p_{t-1}^1)$$

$$=\tfrac{1}{2}(s_t^2+s_{t-1}^2)'[N\,I_{N_2}-L_{N_2xN_2}]^{-1}L_{N_2xN_1}(\ln p_t^1-\ln p_{t-1}^1)$$

$$=\tfrac{1}{2N_1}(s_t^2+s_{t-1}^2)'L_{N_2xN_1}(\ln p_t^1-\ln p_{t-1}^1)$$

$$=\left(\frac{1}{2N_1}\right)\begin{bmatrix}\sum_{i=N_1+1}^{N}(s_{it}+s_{it-1})\\ \vdots \\ \sum_{i=N_1+1}^{N}(s_{it}+s_{it-1})\end{bmatrix}'(\ln p_t^1-\ln p_{t-1}^1).$$

Notice that $\tfrac{1}{2}\left(\sum_{i=N_1+1}^{N}(s_{it}+s_{it-1})\right)$ equals $1-\tfrac{1}{2}\left(\sum_{i=1}^{N_1}s_{it}+s_{it-1}\right)$. Substituting these results into the right-hand side of (A2), we can combine the first and second terms as:

$$\tfrac{1}{2}(s_t^1+s_{t-1}^1)'(\ln p_t^1-\ln p_{t-1}^1)-\tfrac{1}{2}(s_t^2+s_{t-1}^2)'\left[B^{22}\right]^{-1}B^{21}(\ln p_t^1-\ln p_{t-1}^1)$$

$$=\left\{\frac{(s_t^1+s_{t-1}^1)'}{2}+\left(\frac{1}{N_1}\right)\begin{bmatrix}1-\tfrac{1}{2}\left(\sum_{i=1}^{N_1}s_{it}+s_{it-1}\right)\\ \mathrm{M} \\ 1-\tfrac{1}{2}\left(\sum_{i=1}^{N_1}s_{it}+s_{it-1}\right)\end{bmatrix}'\right\}(\ln p_t^1-\ln p_{t-1}^1)$$

$$=\sum_{i=1}^{N_1}\tfrac{1}{2}(\bar{s}_{it}+\bar{s}_{it-1})(\ln p_{it}-\ln p_{it-1}),$$

where,

$$\bar{s}_{i\tau}\equiv s_{i\tau}+\tfrac{1}{N_1}\left(1-\sum_{i=1}^{N_1}s_{i\tau}\right),\ \text{for } i = 1,\ldots,N_1, \text{ and } \tau = t\text{-}1, t.$$

Reintroducing the notation (i,j) to denote each product, and noting that N_1 equals $\overline{N}$ as used in the text, this gives us equation (11).

The final term in (A2) is also simplified using (A3). Substituting for B^{22} and dropping the negative sign for notation convenience, the final term in (A3) becomes:

$$\left(\frac{1}{2\gamma}\right)(s_t^2+s_{t-1}^2)'\left[I_{N_2}-\left(\frac{1}{N}\right)L_{N_2xN_2}\right]^{-1}(s_t^2-s_{t-1}^2)'$$

$$=\left(\frac{1}{2\gamma}\right)(s_t^2+s_{t-1}^2)'\left\{I_{N_2}+\left(\frac{1}{N}\right)L_{N_2xN_2}+\left(\frac{1}{N}\right)^2L^2_{N_2xN_2}+\left(\frac{1}{N}\right)^3L^3_{N_2xN_2}\cdots\right\}(s_t^2-s_{t-1}^2)'$$

$$=\left(\frac{1}{2\gamma}\right)\left\{\sum_{i=N-N_2}^{N}[(s_{it})^2-(s_{it-1})^2]+\left(\frac{1}{N}\right)\left[\left(\sum_{i=N-N_2}^{N}s_{it}\right)^2-\left(\sum_{i=N-N_2}^{N}s_{it-1}\right)^2\right]\left[1+\left(\frac{N_2}{N}\right)+\left(\frac{N_2}{N}\right)^2+\dots\right]\right\}$$

$$=\left(\frac{1}{2\gamma}\right)\left\{\sum_{i=N-N_2}^{N}[(s_{it})^2-(s_{it-1})^2]+\left(\frac{1}{N_1}\right)\left[\left(\sum_{i=N-N_2}^{N}s_{it}\right)^2-\left(\sum_{i=N-N_2}^{N}s_{it-1}\right)^2\right]\right\}.$$

Again reintroducing the notation (i,j) to denote each product, and noting that $N_1=\overline{N}$ and that i= $N–N_2,\dots,N$ are not in the set $\overline{I}$, this gives us equation (12). QED

Proof of Theorem 2:

First, we need to show that (19) is a second-order approximation to (18), around the point where (20) holds and $s_{it}N_t/\gamma(N_t-1)=0$. To this end, express the right of (18) as $\sum_j s^i_{jt}\ln(1+s^i_{jt}x)$, with $x=s_{it}N_t/\gamma(N_t-1)$. We wish to show that the first and second derivatives of this function with respect to s^i_{jt} and x equal the first and second derivatives of $\ln[1+\sum_j(s^i_{jt})^2x]$, evaluated at the point where (20) holds and x = 0. We have:

$$\left.\frac{\partial}{\partial s^i_{jt}}\right|_{\substack{(20),\\x=0}}\sum_j s^i_{jt}\ln(1+s^i_{jt}x)=0=\left.\frac{\partial}{\partial s^i_{jt}}\right|_{\substack{(20),\\x=0}}\ln[1+\sum_j(s^i_{jt})^2x]$$

$$\left.\frac{\partial^2}{\partial s^i_{kt}\partial s^i_{jt}}\right|_{\substack{(20),\\x=0}}\sum_j s^i_{jt}\ln(1+s^i_{jt}x)=0=\left.\frac{\partial^2}{\partial s^i_{kt}\partial s^i_{jt}}\right|_{\substack{(20),\\x=0}}\ln[1+\sum_j(s^i_{jt})^2x]$$

$$\left.\frac{\partial}{\partial x}\right|_{\substack{(20),\\x=0}}\sum_j s^i_{jt}\ln(1+s^i_{jt}x)=\sum_j(s^i_{jt})^2=\left.\frac{\partial}{\partial x}\right|_{\substack{(20),\\x=0}}\ln[1+\sum_j(s^i_{jt})^2x]$$

$$\left.\frac{\partial^2}{\partial s^i_{jt}\partial x}\right|_{\substack{(20),\\ x=0}} \sum\nolimits_j s^i_{jt}\ln(1+s^i_{jt}x) = 2s^i_{jt} = \left.\frac{\partial^2}{\partial s^i_{jt} x}\right|_{\substack{(20),\\ x=0}} \ln[1+\sum\nolimits_j (s^i_{jt})^2 x]$$

$$\left.\frac{\partial^2}{\partial x^2}\right|_{\substack{(20),\\ x=0}} \sum\nolimits_j s^i_{jt}\ln(1+s^i_{jt}x) = -\frac{1}{\bar{N}_i^2} = \left.\frac{\partial^2}{\partial x^2}\right|_{\substack{(20),\\ x=0}} \ln[1+\sum\nolimits_j (s^i_{jt})^2 x],$$

where we note that the summations above are over $j\in \bar{J}_i$ and only the last line relies on $s^i_{jt} = (1/\bar{N}_i)$, from (20).

Then using (20), we replace $s_{ij\tau}$ in (11) by $s_{it}/\bar{N}_i$, and use this in (10) to obtain:

$$\ln\left(\frac{e_t}{e_{t-1}}\right) = V + \sum_{i\in\bar{I}} \tfrac{1}{2}(s_{it}+s_{it-1}) \sum_{j\in\bar{J}_i} \tfrac{1}{\bar{N}_i}(\ln p_{ijt} - \ln p_{ijt-1}) +$$

$$\frac{1}{\bar{N}}\left(1 - \tfrac{1}{2}\sum_{i\in\bar{I}}\sum_{j\in\bar{J}_i}\frac{s_{it}}{\bar{N}_i} - \tfrac{1}{2}\sum_{i\in\bar{I}}\sum_{j\in\bar{J}_i}\frac{s_{it-1}}{\bar{N}_i}\right)\sum_{i\in\bar{I}}\sum_{j\in\bar{J}_i}(\ln p_{ijt} - \ln p_{ijt-1})$$

$$= V + \sum_{i\in\bar{I}} \tfrac{1}{2}(s_{it}+s_{it-1})(\overline{\ln p_{it}} - \overline{\ln p_{it-1}}) + \frac{1}{\bar{N}}\left(1 - \tfrac{1}{2}\sum_{i\in\bar{I}} s_{it} - \tfrac{1}{2}\sum_{i\in\bar{I}} s_{it-1}\right)\sum_{i\in\bar{I}} \bar{N}_i\ (\overline{\ln p_{it}} - \overline{\ln p_{it-1}}),$$

where $\overline{\ln p_{it}} \equiv \frac{1}{\bar{N}_i}\sum_{j\in\bar{J}_i} \ln p_{ijt}$ is the unweighted mean of the log-prices for country i. Again from (20), these are identical to the weighted mean of log-prices defined in the text, $\ln p_{it} \equiv \sum_{j\in\bar{J}_i} s^i_{jt}\ln p_{ijt}$. Then using the shares in Theorem 2 and (19), we re-write the above result as:

$$\ln\left(\frac{e_t}{e_{t-1}}\right) = V + \sum_{i\in\bar{I}} \tfrac{1}{2}(\bar{s}_{it}+\bar{s}_{it-1})(\ln p_{it} - \ln p_{it-1}) = V + \sum_{i\in\bar{I}} \tfrac{1}{2}(\bar{s}_{it}+\bar{s}_{it-1})(\ln C'_{it} - \ln C'_{it-1}) + P,$$

with P defined as in (21). Then using $\ln(1+x)\approx x$, P can be re-written as:

$$\sum_{i\in\bar{I}} \tfrac{1}{2}(\bar{s}_{it}+\bar{s}_{it-1})\left\{\ln\left[1+\frac{H_{it}s_{it}N_t}{\gamma(N_t-1)}\right] - \ln\left[1+\frac{H_{it-1}s_{it-1}N_{t-1}}{\gamma(N_{t-1}-1)}\right]\right\}$$

$$\approx \sum_{i\in\bar{I}} \tfrac{1}{2}(\bar{s}_{it}+\bar{s}_{it-1})\left[\frac{H_{it}s_{it}N_t}{\gamma(N_t-1)} - \frac{H_{it-1}s_{it-1}N_{t-1}}{\gamma(N_{t-1}-1)}\right]$$

$$= \frac{1}{2\gamma}\sum_{i\in \bar{I}}(\bar{s}_{it} + \bar{s}_{it-1})\left[(H_{it}s_{it} - H_{it-1}s_{it-1}) + \frac{H_{it}s_{it}}{(N_t - 1)} - \frac{H_{it-1}s_{it-1}}{(N_{t-1} - 1)}\right]$$

Using the formula for the defined shares in Theorem 2, we can re-write P as:

$$P = \frac{1}{2\gamma}\sum_{i\in \bar{I}}\left[(s_{it} + s_{it-1}) + \frac{\bar{N}_i}{\bar{\bar{N}}}\left(\sum_{i\notin \bar{I}} s_{it} + \sum_{i\notin \bar{I}} s_{it-1}\right)\right](H_{it}s_{it} - H_{it-1}s_{it-1})$$

$$+ \frac{1}{2\gamma}\sum_{i\in \bar{I}}(\bar{s}_{it} + \bar{s}_{it-1})\left[\frac{H_{it}s_{it}}{(N_t - 1)} - \frac{H_{it-1}s_{it-1}}{(N_{t-1} - 1)}\right]$$

$$= \frac{1}{2\gamma}\sum_{i\in \bar{I}}(s_{it} + s_{it-1})(H_{it}s_{it} - H_{it-1}s_{it-1}) + \frac{1}{2\gamma}\left(\sum_{i\notin \bar{I}} s_{it} + \sum_{i\notin \bar{I}} s_{it-1}\right)\sum_{i\in \bar{I}}\frac{\bar{N}_i}{\bar{\bar{N}}}(H_{it}s_{it} - H_{it-1}s_{it-1})$$

$$+ \frac{1}{2\gamma}\sum_{i\in \bar{I}}(\bar{s}_{it} + \bar{s}_{it-1})\left[\frac{H_{it}s_{it}}{(N_t - 1)} - \frac{H_{it-1}s_{it-1}}{(N_{t-1} - 1)}\right].$$

From (20) note that $H_{it} = H_{it-1} = 1/\bar{N}_i$ for $i \in \bar{I}$, and using this repeatedly we can simplify P as:

$$P = \frac{1}{2\gamma}\sum_{i\in \bar{I}}(H_{it}s_{it}^2 - H_{it-1}s_{it-1}^2) - \frac{1}{2\gamma\bar{\bar{N}}}\left(\sum_{i\notin \bar{I}} s_{it} + \sum_{i\notin \bar{I}} s_{it-1}\right)\left(\sum_{i\notin \bar{I}} s_{it} - \sum_{i\notin \bar{I}} s_{it-1}\right)$$

$$+ \frac{1}{2\gamma}\sum_{i\in \bar{I}}(\bar{s}_{it} + \bar{s}_{it-1})\left[\frac{H_{it}s_{it}}{(N_t - 1)} - \frac{H_{it-1}s_{it-1}}{(N_{t-1} - 1)}\right]$$

Then substituting for V from (12), we obtain the result shown in (22). QED

Appendix B: Data

The dataset used for this project contains quantity, value, and price information aggregated at the HS-10 digit level, as well as HS-4 digit level Herfindahl Indexes, for the U.S. and all countries exporting to the U.S. for every year from 1992 to 2005.

One challenge in piecing together this dataset was calculating the amount of U.S. absorption produced in the U.S. We begin with the identity that the U.S. supply of U.S. absorption is equal to the difference between U.S. production and exports. We obtained data on industry-level production from the Bureau of Economic Analysis at www.bea.gov and export data from http://www.internationaldata.org/. Unfortunately, the BEA production data are classified according to the SIC system for years 1992 to 1997 and according to the NAICS system for years 1998 to 2005, while the trade data is at the HS-10 digit level. Addressing this complication required a two-step process: the first step was to adjust the BEA production data so that the data are on the NAICS level for all years within the sample. The second step was to use our import/export data (containing both NAICS and HS-10 digit codes) and our newly created NAICS level production data to infer domestically produced absorption at the HS-10 digit level, as described below.

It is not easy to concord SIC and NAICS categories because there is not always a one-to-one mapping between the two. To deal with this issue, we first used a NAICS-SIC concordance from the BEA to convert the SIC data to the NAICS level. The absence of a one-to-one mapping meant that sometimes we would observe large jumps in a NAICS category derived from SIC data from 1997 relative to what the 1998. In order to deal with this problem, we used a "bridging dataset", from the U.S. Department of Commerce, containing SIC level values for both 1997 and 1998. This enabled us to construct a ratio between the actual NAICS output levels and the NAICS levels that we constructed from the SIC data for 1998. We then multiplied all of NAICS data that was constructed from the SIC data by this ratio. If a SIC sector did not match any NAICS sector we dropped the observations prior to 1998 in the estimation. We also dropped all changes between 1997 and 1998 in the regressions where we estimated γ, so that concordance problems would not affect our estimates.

After our BEA data was brought to the NAICS level, we use it, along with our import and export data, to calculate HS-10 digit level U.S. domestic supply. We begin with the identity that U.S. supply for the domestic market at the NAICS level – denoted by k – equals U.S. production at the NAICS level less U.S. exports:

$$\text{Supply}_t^k = \text{Production}_t^k - \text{Exports}_t^k .$$

Using the NAICS import data, we can compute the share of U.S. supply in apparent consumption according to the following formula:

$$\text{Share}_t^k = \frac{\text{Supply}_t^k}{\text{Supply}_t^k + \text{Imports}_t^k}$$

By assuming that the U.S. share of a NAICS code is equal to that of the U.S. share in a corresponding HS-10 code, we calculate supply at the HS-10 digit level using the following formula:

$$\text{Share}_t^k = \text{Share}_t^{\text{HS10}} \quad \text{and} \quad \text{Share}_t^{\text{HS10}} = \frac{\text{Supply}_t^{\text{HS10}}}{\text{Supply}_t^{\text{HS10}} + \text{Imports}_t^{\text{HS10}}}$$

$$\Rightarrow \text{Supply}_t^{\text{HS10}} = \frac{\text{Share}_t^k}{\left(1 - \text{Share}_t^k\right)} \text{Imports}_t^{\text{HS10}} .$$

We next needed to merge in data for Herfindahl indexes for domestic firms and exporters to the U.S. For land shipments from Canada, we purchased Herfindahl indexes at the 4-digit

Harmonized system (HS) level, for 1996 and 2005, from Statistics Canada. These Canadian Herfindahl indexes were constructed from firm-level export data to the U.S.

For land shipments from Mexico, the Herfindahl indexes were constructed using data sourced from the *Encuesta Industrial Anual* (Annual Industrial Survey) of the *Instituto Nacional de Estadistica y Geografia*. This data contains firm-level exports for 205 CMAP94 categories for 1993 and 2003. We also obtained the export Herfindahl for 232 categories at the HS-4 level. These categories cover the most important Mexican export sectors.

For all other major exporters to the U.S., we computed these Herfindahls for *sea* shipments from PIERS (www.piers.com), for 1992 and 2005. PIERS collects data from the bill of landing for every container that enters a U.S. port. The median country exports about 80 percent of its goods by sea. Thus for the typical country in our sample, the sea data covers a large fraction of their exports. Although purchasing the disaggregated data is prohibitively expensive, we were able to obtain information on shipments to the U.S. for the 50,000 largest exporters to the U.S., for 1992 and 2005. For each exporter and year, we obtained the estimated value, quantity and country of origin of the top five HS-4 digit sectors in which the firm was active. We also obtained this data for the top ten HS-4 digit sectors for the largest 250 firms in each year.

The Piers data has a number of limitations relative to other firm level data sets. The first is relatively minor: we do not have the universe of exporters but only the largest ones. This turns out not to be a serious problem because the aggregate value of these exporters is typically within 5 percent of total sea shipments. Thus, smaller exporters are unlikely to have a qualitatively important impact on our results.

A larger problem is that the PIERS data only comprises sea shipments and thus we have no information in these data on land and air shipments. This means that we have to adjust our

Herfindahl indexes to take into account land and air shipments. The Herfindahl of country i's exports in sector k can be written as

$$H_{it}^{k} = H_{it}^{kSea}\left(\frac{V_{it}^{kSea}}{V_{it}^{kTotal}}\right)^2 + H_{it}^{kNon-Sea}\left(1-\frac{V_{it}^{kSea}}{V_{it}^{kTotal}}\right)^2, \qquad \text{(B1)}$$

where V_{it}^{kSea} $\left(V_{it}^{kTotal}\right)$ denotes the value of sea (total) shipments and $H_{it}^{kNon-Sea}$ is the Herfindahl for non-sea exporters, which is defined analogously as the sea Herfindahl. We do not have a measure of $H_{it}^{kNon-Sea}$, but theory does place bounds on the size of the Herfindahl since the true index must be contained in the following set, obtained with $H_{it}^{kNon-Sea}$ =1 or 0:

$$\left[H_{it}^{kSea}\left(\frac{V_{it}^{kSea}}{V_{it}^{Total}}\right)^2, H_{it}^{kSea}\left(\frac{V_{it}^{kSea}}{V_{it}^{Total}}\right)^2 + \left(1-\frac{V_{it}^{kSea}}{V_{it}^{kTotal}}\right)^2\right].$$

For most sectors the share of sea shipments in total shipments is quite high, so these bounds are quite tight. In the analysis we assume that $H_{it}^{kSea} = H_{it}^{kNon-Sea}$, but our results do not change qualitatively if we assume that $H_{it}^{kNon-Sea}$ or 0.[17]

For the U.S. Herfindahls, we rely on data from the Census of Manufactures are at the NAICS 6-digit level. Unfortunately, this is more aggregate than the 4-digit HS level at which we have the foreign export Herfindahl indexes. Accordingly, we need to convert the U.S. Herfindahl indexes from the NAICS 6-digit level to the HS 4-digit level. Slightly abusing our earlier country notation, let $i \in I_k$ denote a 4-digit sector i *within* the NAICS code k. Then the Herfindahl for 4-digit sector i is $H_{it}^{k} \equiv \sum_{j\in J_i}(s_{jt}^{i})^2$, where s_{jt}^{i} is the share of firm $j \in J_i$ in sector i. We see that the overall Herfindahl in NAICS code k is:

[17] One can see this from a simple example. Our median sea Herfindahl is 0.6 and our median share of sea shipments is 0.8. This means that the true Herfindahl ranges from .38 to .42 and our estimate would be 0.41. Nevertheless, we are implicitly assuming that goods shipped by air and goods shipped by sea are not the same. We justify this assumption because it costs substantially more to ship goods by air, and thus the mode of shipment is likely to differentiate the goods in some important ways.

$$\sum_{i\in I_k} H_{it}^k \left(s_{it}^k\right)^2 = \sum_{i\in I_k}\sum_{j\in J_i}(s_{jt}^i)^2(s_{it}^k)^2 = \sum_{j\in J_k}(s_{jt}^k)^2 \equiv H_t^k\ , \tag{B2}$$

where s_{it}^k is the share of 4-digit HS sector i within NAICS sector k, and $s_{jt}^k = s_{jt}^i s_{it}^k$ is the share of product j within the NAICS sector, $j\in J_k$. In words, the inner-product of the Herfindahl firm indexes and the squared sector shares, on the left of (B2) is exactly the right way to aggregate these indexes to obtain an *overall Herfindahl for the good k in question,* on the right of (B2). One of the problems that we faced is that we know H_t^k but not H_{it}^k. A solution can be obtained by assuming that H_{it}^k is equal across all 4-digit sectors $i\in k$, in which case we solve for H_{it}^k as:

$$H_{it}^k = H_t^k \,/ \sum_{i\in I_k}(s_{it}^k)^2\ . \tag{B3}$$

In other words, the 4-digit HS Herfindahl is estimated by dividing the 6-digit NAICS Herfindahl by the corresponding Herfindahl index of 4-digit HS shares within the 6-digit sector. This simple solution assumes that the 4-digit HS Herfindahl indexes are constant within a sector, but is the best that we can do in the absence of additional data.

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Table 1

Ranking in Terms of Share of U.S. Total Absorption

1992				1997			
Country	Herfindahl Index	Share	Weighted Ave. $H_{it}S_{it}$	Country	Herfindahl Index	Share	Weighted Ave. $H_{it}S_{it}$
United States	0.147	0.801	0.1114	United States	0.155	0.745	0.1107
Canada	0.245	0.038	0.0106	Canada	0.252	0.052	0.0132
Japan	0.310	0.036	0.0094	Japan	0.313	0.035	0.0100
Mexico	0.393	0.012	0.0040	Mexico	0.407	0.024	0.0086
German	0.358	0.010	0.0030	China	0.293	0.017	0.0019
China	0.366	0.010	0.0011	German	0.357	0.012	0.0035
Taiwan	0.365	0.008	0.0015	United Kingdom	0.331	0.009	0.0028
South Korea	0.419	0.007	0.0017	Taiwan	0.369	0.008	0.0022
United Kingdom	0.309	0.007	0.0020	South Korea	0.396	0.007	0.0023
Saudi Arabia	0.427	0.005	0.0010	Malaysia	0.398	0.006	0.0016
Weighted Ave.	0.160		0.0781	Weighted Ave.	0.170		0.0692

1998				2005			
Country	Herfindahl Index	Share	Weighted Ave. $H_{it}S_{it}$	Country	Herfindahl Index	Share	Weighted Ave. $H_{it}S_{it}$
United States	0.183	0.781	0.1392	United States	0.189	0.692	0.1289
Canada	0.249	0.043	0.0111	Canada	0.242	0.056	0.0146
Japan	0.318	0.030	0.0085	China	0.188	0.041	0.0026
Mexico	0.419	0.022	0.0083	Mexico	0.403	0.031	0.0101
China	0.280	0.017	0.0017	Japan	0.331	0.025	0.0078
German	0.332	0.012	0.0034	German	0.335	0.015	0.0049
United Kingdom	0.331	0.007	0.0025	United Kingdom	0.331	0.009	0.0025
Taiwan	0.340	0.007	0.0018	South Korea	0.338	0.009	0.0028
South Korea	0.377	0.006	0.0020	Venezuala	0.556	0.008	0.0046
France	0.371	0.005	0.0020	Saudi Arabia	0.447	0.006	0.0019
Weighted Ave.	0.190		0.0903	Weighted Ave.	0.191		0.0714

Notes:
The Herfindahl Index is the weighted average of the country's Herfindahl Index, where the weights correspond to the share of each HS-4 sector in U.S. apparent consumption. "Share" s_{it} is defined to be the country's share of U.S. apparent consumption. The "Weighted Average $H_{it}s_{it}$" is the weighted average of the Herfindahl Index in sector i in year t multiplied by that country's share of U.S. apparent consumption; the weights are the same as before. The last row reports a weighted average across all countries using each country's share of U.S. apparent consumption as weights. Thus, the number shown in the last row and column of each panel is $\Sigma_i H_{it}(s_{it})^2$, averaged across sectors.

Table 2

Distribution of γ Estimates

Statistic	Value	Standard Deviation
Mean	11.90	1.75
Median	0.19	0.01
Median Number of Varieties per HS4	94	

Table 3

Gamma Values From Sectors with High Shares of Domestic Absorption

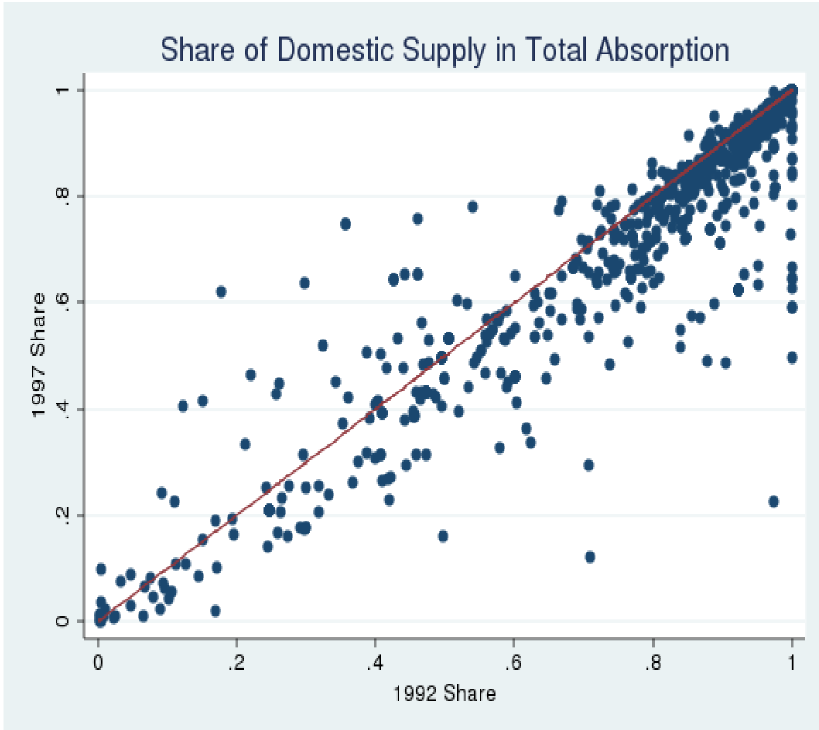
Hs4	γ	Average Share of Total Absorption
Passenger motor vehicles	0.14	0.07
Parts and accessories for non-passenger motor vehicles	0.39	0.05
Crude petroleum	0.76	0.04
Automatic data processing machines	0.18	0.03
Non-military aircrafts	0.06	0.02
Cartons, boxes, cases, bags and other packing containers	0.25	0.02
Cell phones	0.07	0.01
Cigarettes	1.41	0.01
Plastics	0.05	0.01
Natural Gas	1.11	0.01

Table 4

Partial Markup and Variety Effects

Range of Herfindahl Movement Defined As "Common"	Specification	P	V	Total (P+V)
(1/1.3,1.3)	Benchmark	-0.017	-0.037	-0.054
(1/1.3,1.3)	No passenger vehicles sector	-0.013	-0.013	-0.026
(1/1.1,1.1)	Change Herfindahl range	-0.015	-0.030	-0.044
(1/1.2,1.2)	Change Herfindahl range	-0.016	-0.035	-0.050
(1/1.4,1.4)	Change Herfindahl range	-0.014	-0.035	-0.049
(1/1.5,1.5)	Change Herfindahl range	-0.010	-0.029	-0.039
(1/1.3,1.3)	Sum of 92-97 and 98-05	-0.018	-0.034	-0.051
(1/1.3,1.3)	Herfindahls Set to 1992 Values	-0.032	-0.004	-0.036

Figure 1


Share of Domestic Supply in Total Absorption
1997 Share
0
.2
.4
.6
.8
1
1992 Share

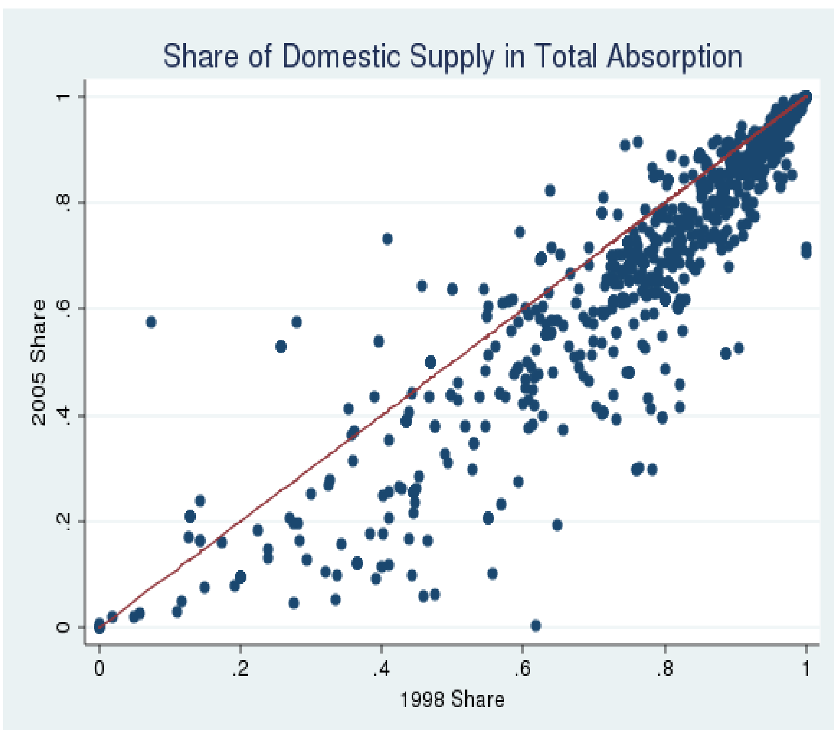

Share of Domestic Supply in Total Absorption
2005 Share
0
.2
.4
.6
.8
1
1998 Share

Figure 2

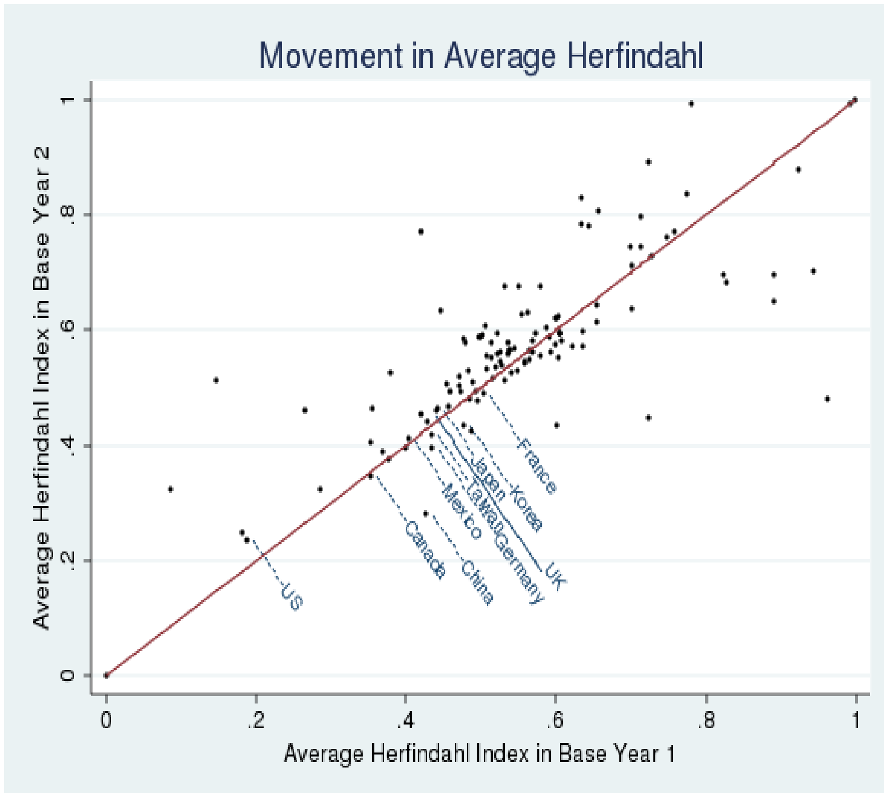


Figure 3a: Distribution of 2γV

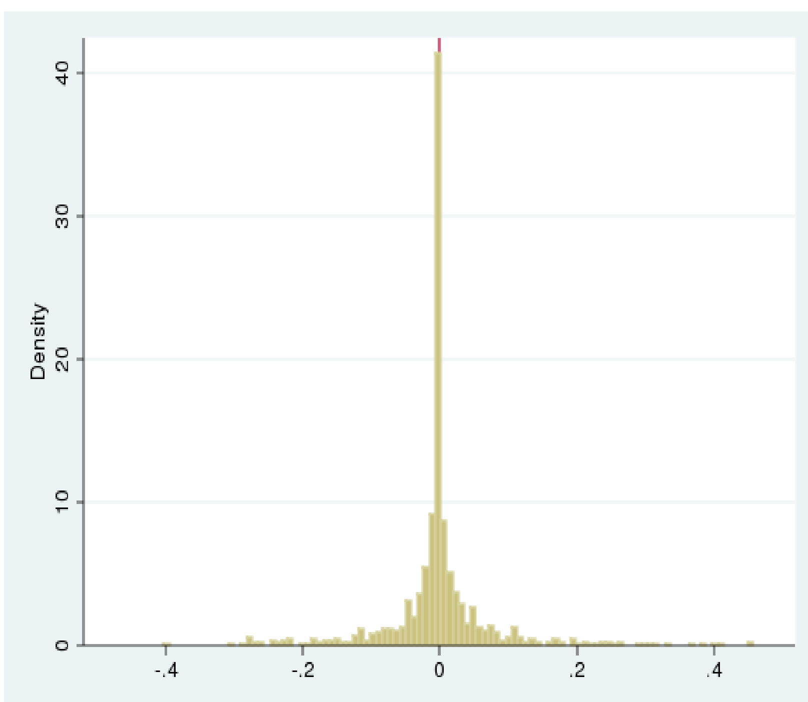


Figure 3b:Distribution of 2γP

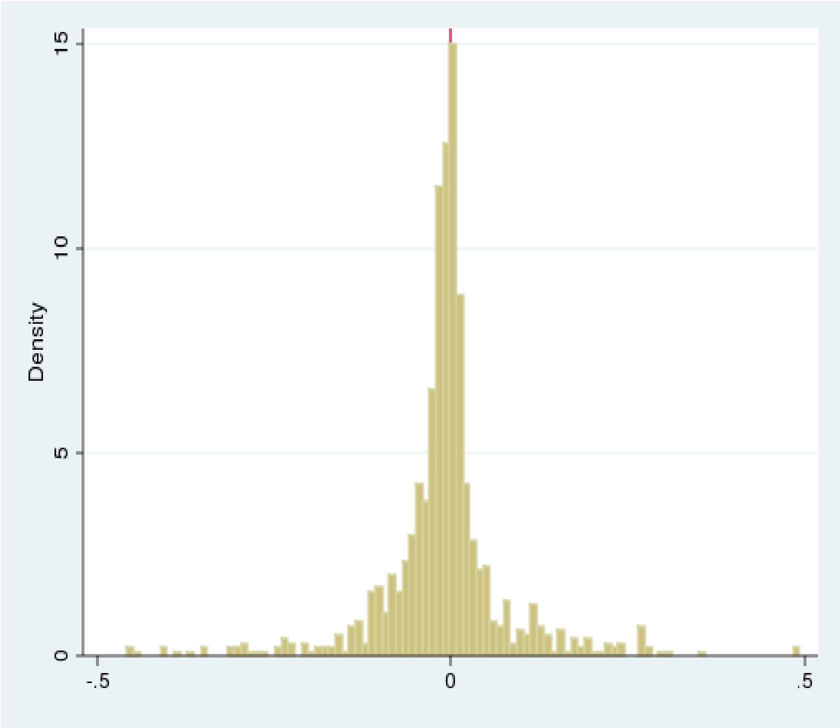


Figure 3c: Distribution of 2γ(P-V)