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FOOD PROTECTION FOR SALE

Abstract

This article tests the Protection for Sale (PFS) model using detailed data from U.S. food processing industries from 1978 to 1992 under alternative import demand specifications. All empirical results support the PFS model predictions and previous empirical work qualitatively. Although welfare weights are very sensitive to import demand specification, a surprising result is that we obtain weights between 2.6 and 3.6 for domestic welfare using import slopes or elasticities derived from domestic demand and supply functions. In contrast, results based on import slopes or elasticities from directly specified import demands (including the Armington model) yield the usual, unrealistically large estimates for the domestic welfare weight. We contend that the latter empirical paradox arises mainly because the explanatory variables tend to be extremely large for industries with low import ratios and/or low estimated elasticities or slopes resulting from relatively volatile import prices. The results with derived import parameters point to a much stronger role of campaign contributions within the PFS model than previously found. They also suggest that the commonly-used Armington estimates may not be appropriate for estimating the PFS model.

Key words: Trade protection, tariffs, lobbying, political economy, food manufacturing

JEL codes: F13, F1, L66, C12

FOOD PROTECTION FOR SALE

Introduction

The most influential of the last wave of the political economy of trade protection models is the "Protection for Sale" (PFS) model developed by Grossman and Helpman (1994, henceforth GH). Several studies have confirmed its qualitative predictions (e.g., Goldberg and Maggi, 1999; Gawande and Bandyopadhyay, 2000; Eicher and Osang, 2002; Matschke and Sherlund, 2004), but all have obtained unrealistically large estimates of the weight the government places on general welfare vs. the weight on campaign contributions, leading to the conclusion that protection is not for sale.[1] These large general welfare weight estimates also create a cognitive dissonance between the typical levels of campaign contributions by industries and the much larger magnitude of trade policy benefits they receive, questioning the truthful contribution assumption maintained in the PFS model (Lopez, 2001).

This article applies the PFS model to a sample of U.S. food processing industries. These industries provide a good case study to analyze trade protection. First, trade protection varies substantially across industries, from those receiving little or no protection (e.g., roasted coffee and macaroni and spaghetti) to those with nominal protection coefficients exceeding 50% (e.g., cane sugar, dairy products, and frozen specialties). Second, import penetration ratios range from less than 2% for milk to over 40% for wine and spirits. Third, these industries show wide variation in political participation and organization, as reflected by their campaign contributions. Fourth, they constitute the largest manufacturing sector in the U.S. economy in terms of value of shipments.[2]

Focusing on food processing industries allows us to look more closely at the determinants of trade protection. First, all previous empirical studies of the PFS model take estimates of import demand elasticities from outside sources (for the U.S., the commonly used source is the study by Shiells, Stern, and Deardorff (1986), which provides elasticity estimates at the 3-digit SIC level). In contrast, we derive 4-digit SIC import demand elasticity and slope estimates directly from the data used in our sample.[3] Second, all previous PFS studies either abstract from the existence of intermediate goods (e.g. Goldberg and Maggi, 1999) or assume just one intermediate good that is freely traded and used by all industries within the sample (Gawande and Bandyopadhyay, 2000). In contrast, we explicitly model the fact that food processing industries buy products from each other and as buyers may actually lobby for lower trade protection for input-providing industries. Third, we use tariff rates and tariff equivalents to measure trade protection, whereas the previous literature exclusively uses NTB coverage ratios for the United States. The PFS model, however, does not provide any predictions for NTB coverage ratios.

The empirical results show that the estimated weights are quite sensitive to the precise import demand specification and that the weights are much smaller (between 2.6 and 3.6) when they are derived from estimated domestic demand and supply. Using import slopes or elasticities based on direct import demand specifications (including the Armington (1969) model) yields the usual high weights placed on domestic welfare found in the previous literature, reinforcing the conclusion that protection is not for sale. The latter result stems in part from observations for industries with low import penetration ratios and possibly low values for import elasticities and/or slopes due to a much larger variation in import prices than domestic ones. Thus, the results with derived import demand slopes or elasticities indicate that the role of campaign contributions within the PFS may be much stronger than previously found.

The Protection for Sale Model

In the PFS model (summarized here for exposition purposes), the government values both the total level of political contributions and the aggregate well-being of the population so that the equilibrium tariff vector maximizes governmental welfare:

$$aW^{G} + \sum_{j \in L} C_{j}\ , \qquad (1)$$

where a is the weight given to general welfare W^{G}, L denotes the set of politically organized sectors, and C_{j} is contributions by sector *j*. Letting W^{k} denote the welfare of a specific sector *k*, the sum of governmental welfare and welfare of lobby *k* is given by:

$$aW^{G} + W^{k} + \sum_{j \in L, j \neq k} C_{j}\ , \qquad (2)$$

for all $k \in L$. The first-order condition w.r.t. the specific tariff t_{i} for maximization of governmental welfare yields:

$$a\frac{\partial W^{G}}{\partial t_{i}} + \sum_{j \in L} \frac{\partial C_{j}}{\partial t_{i}} = 0\ . \qquad (3)$$

The first-order condition w.r.t. t_{i} for maximization of governmental welfare and welfare of lobby k yields:

$$a\frac{\partial W^{G}}{\partial t_{i}} + \frac{\partial W^{k}}{\partial t_{i}} + \sum_{j \in L, j \neq k} \frac{\partial C_{j}}{\partial t_{i}} = 0 \qquad (4)$$

From equations (3) and (4), local truthfulness easily follows for organized sectors:

$$\frac{\partial W^{k}}{\partial t_{i}} = \frac{\partial C_{k}}{\partial t_{i}}\ . \qquad (5)$$

GH show that under global truthfulness, a lobby compensates the government for the domestic welfare loss that arises from its lobbying, i.e. the government is as well off as if the lobby did not exist.

In the following, we assume that industries lobby for trade policy to increase their profits, but that lobbies only represent a small fraction of the population and as such do not take into account the effects of trade protection on consumer surplus from consumption of final goods and on tariff revenue. This is not to say, however, that the lobby of industry i only cares about the tariff t_i on its own good. Lobby i is also concerned about the prices of its inputs and will lobby for negative protection for its input goods. Therefore, we have

$$\frac{\partial W^i}{\partial t_i} = Q_i, \tag{6}$$

where Q_i denotes the output of industry i, and

$$\frac{\partial W^k}{\partial t_i} = -X_i^k, \tag{7}$$

where X_i^k denotes the input of good i in industry k and $k \neq i$.

Summing equation (4) over all lobbies, we obtain

$$ma\frac{\partial W^G}{\partial t_i} + (m-1)\sum_{j \in L} \frac{\partial C_j}{\partial t_i} + \sum_{j \in L} \frac{\partial W^j}{\partial t_i} = 0, \tag{8}$$

where m denotes the number of lobbies. Substituting from (3), we obtain

$$a\frac{\partial W^G}{\partial t_i} + \sum_{j \in L} \frac{\partial W^j}{\partial t_i} = 0, \tag{9}$$

where

$$\frac{\partial W^G}{\partial t_i} = Q_i - D_i + M_i + t_i M_i^{'} = t_i M_i^{'}, \qquad (10)$$

where D_i stands for the entire consumption of good i. The equilibrium specific tariff is then given by

$$t_i = \frac{1}{a}\frac{I_i Q_i - \sum_j I_j X_i^j}{M_i^{'}}, \qquad (11)$$

or, after rewriting,

$$\tau_i = \frac{t_i}{p_i} = \frac{t_i^a}{1+t_i^a} = \frac{1}{a}\frac{I_i Q_i - \sum_j I_j X_i^j}{M_i e_i}, \qquad (12)$$

where I_i is the indicator variable for lobbying by industry i, t_i^a denotes ad-valorem tariff rates, e_i is the absolute value of the price elasticity of imports, p_i is the domestic price, and other notation is as defined before. Taking buyer lobbying into account will increase the estimate of $\frac{1}{a}$ and decrease the estimate of a.

From (12), the GH model yields three behavioral predictions to be tested:[4] (1) industries that are not politically organized face lower rates of protection than those that are organized; (2) industries that face organized opposition from buyers are granted lower levels of protection; and (3) for protected industries with a constant share of shipments that go to organized buyers, the level of protection is inversely related to the price elasticity of imports and to import penetration.

Empirical Implementation

Equations (11) and (12) provide the conceptual basis for the empirical models to be estimated. The data set to be used contains a number of industries over a number of years. Thus, denote industries by the subscript i and years by the subscript t (not to be confused with the specific tariff notation).

We estimate import slopes and elasticities for our data set for all industries in the sample, taking advantage of the time variation in prices and imports. Since import slopes $M_i^{'}$ (and price elasticities of imports for that matter) have to be estimated, it is instructive to follow Goldberg and Maggi's approach to deal with errors in estimates and pass the import slope to the left-hand side so that the estimating equation becomes

$$T_{it}^{M} = t_{it}\hat{M}_i^{'} = \beta Z_{it}^{M} + \varepsilon_{it}^{M}, \qquad (13)$$

where $\hat{M}_i^{'}$ is an estimate of the import slope expressed in absolute value, $\beta = 1/a$, $Z_{it}^{M} = I_{it}Q_{it} - \sum_j I_{jt}X_{it}^{j}$, where j superscripts indicate the political organization and quantity purchases of industry j, and ε_{it}^{M} is an error term. Likewise, (12) can be rewritten as

$$T_{it}^{e} = \tau_{it}\hat{e}_i = \beta Z_{it}^{e} + \varepsilon_{it}^{e}, \qquad (14)$$

where $Z_{it}^{e} = Z_{it}^{M} / M_{it}$.

To operationalize (13) and (14), annual time series data (1978-92) from 24 food processing industries at the 4-digit 1972 SIC level are used, resulting in 360 observations.[5] The domestic input and output values as well as corresponding price indices are taken from the NBER database on manufacturing productivity by Bartelsman and Gray (1996). Output and input quantity indices were obtained by dividing the value of shipments and input expenditures by their respective price indices. The amount of output bought by other food processors was obtained from the 1977, 1982, 1987 and

1992 Benchmark Input-Output Accounts of the United States (U.S. Department of Commerce, various years).[6]

The values of imports at the 4-digit SIC level were taken from Feenstra (1996). Average tariff rates were computed by dividing total duties collected by CIF import values from a tape supplied by the US International Trade Commission (1978-90) and its website (dataweb.usitc.gov) for 1991-92. Tariff-rate equivalents were used for four industries protected by import quotas: sugar (SIC 2061), meat packing (SIC 2011), cheese (SIC 2021), and milk (SIC 2026). The tariff-rate equivalents were taken from two reports of the U.S. International Trade Commission (1990a, 1990b) and a U.S. Department of Agriculture (1994) report.[7]

Data on import prices at the 4-digit SIC level are not readily available. However, the FAO website and Foreign Agricultural Trade of the United States (USDA, various years) databases provided data on quantity and price for most processed agricultural products. Import price indices were constructed by aggregating products by SIC definitions and by weighting available quantity and price values.[8]

Two sets of estimates for import slopes $M_i^{'}$ are obtained: derived slopes from domestic linear supply and demand functions and direct slopes from a linear import demand equation. To obtain derived slopes, a simultaneous equation system of linear domestic demand and supply functions is estimated for each industry via three-stage least squares. The derived slope is obtained as $\hat{M}_i^{'} = \hat{\alpha}_i - \hat{\gamma}_i$, where $\hat{\alpha}_i$ is the estimated domestic demand slope and $\hat{\gamma}_i$ is the estimated domestic supply slope. To obtain direct import slopes, a linear import demand function was estimated via two-stage least squares. Apart from domestic price (for the import demand derived from the demand-supply system) and import price (for direct import demand), the regressors included the domestic output price index, price of raw materials, wage rates, total factor productivity growth, aggregate consumer expenditures, and time index. All nominal prices and consumer expenditures were deflated by the consumer price index. Both sets of import slopes are reported in Table 2.

Three sets of estimates for elasticities e_i are obtained: derived and direct as well as Armington elasticities. As in the case of import slopes, the derived elasticities are estimated using the same set of explanatory variables but applied to a double-log functional form. To keep the import demand elasticities constant over time in the supply-demand framework, supply and demand elasticities are weighted by the industry-specific mean ratios of supply to imports and demand to imports, respectively.[9] The direct elasticities are estimated from a double-log import demand function via two-stage least squares and the same set of explanatory variables. All three elasticity estimates are presented in Table 2.

The Armington estimates are within the range of previous estimates for food manufacturing at the 4-digit SIC level, with an average (absolute) value of 1.265. Lopez and Pagoulatos (2002) estimated an average elasticity of 1.59 for 40 food industries while Gallaway, McDaniel and Rivera (2003) obtained an average of 0.931 for 35 food industries and Reinert and Clinton (1991) obtained an average elasticity of 0.582 for 17 food industries. The average direct import demand elasticity is at 1.458, somewhat higher than average for the Armington elasticities but still within the range of previous estimates at the 4-digit SIC level.

To correct for endogeneity, we employ a two-stage least squares procedure: In the first stage, we regress outputs and imports, respectively, on the exogeneous variables of the domestic market model used to compute import slopes. Thus, our instruments include the price of raw materials, wage rates, total factor productivity growth, aggregate consumer expenditures, and a time index. These instruments are deemed satisfactory for further analysis based on the high correlation between observed and predicted values of outputs and imports.[10] In the second stage, we then replace home-bound industry outputs and imports by their fitted values.

Following Goldberg and Maggi (1999), political action committee (PAC) campaign contributions to congressional candidates were used to construct the political organization variable I_{it}. PAC contributions were first assigned to 4-digit SIC codes for each industry between 1978 and 1992. The PAC data came from bi-annual reports of the Federal Election Commission encompassing the congressional election cycles. Contributions were then deflated by the producer price index (1992 = 1). Estimation proceeded in three steps. First, increments of thresholds of PAC contributions (from \$5,000 to \$200,000 in \$5,000 increments) were used to define I_{it}. Second, a logit model was estimated and those observations with predicted values greater than 0.5 were taken to correspond to organized sectors ($I_{it} = 1$; 0 otherwise).[11] Third, a preliminary version of the PFS model was estimated based on steps 1 and 2 and the t-values and stability of results compared. Robust results were obtained at a threshold PAC contribution of \$10,000.

To assess the implications of using various import demand approaches, five alternative models are estimated and presented: two based on equation (13) with alternative import slopes and three based on (14) with alternative import elasticity estimates. At this juncture, it should be noted that the error terms in (13) and (14) are heteroskedastic. Thus, the equations are estimated using two-stage least squares with heteroskedasticity-robust standard errors.[12]

Data pooling tests were performed three ways: time pooling, industry pooling, and complete pooling. A Chow test of these effects failed to reject time pooling but not industry pooling at the 5% level for all five versions of the PFS model.[13] Given these test results and since we are interested in the structure of protection rather than individual time and industry effects, only the main results with complete pooling are presented.[14]

Empirical Results

Table 3 presents alternative estimates for the PFS models. The β coefficients are statistically significant at the 99% level across all models. The results provide further support for the fundamental predictions of the GH model. Organized sectors receive more protection than unorganized ones, and for organized sectors, protection decreases with import penetration or the price elasticity of import demand. In addition, industries that sell less output to organized buyers receive more protection.

Conceptually, the empirical estimates of equation (14) should yield the same estimate for β using any of the three elasticity estimates and the same relative welfare weight on general welfare as equation (13) using either set of import slope estimates. Empirically, this is not the case. Although all import demand estimates are based on the same data, it is of interest to compare the results, as done below.

While understandably the estimated domestic welfare weights are quite sensitive to import demand specification, ranging from approximately 2.6 to 3,360, some patterns are clear. First, the results based on directly specified import demand elasticities yield the usual, unrealistically high weights on domestic welfare found in previous studies, suggesting that protection is not for sale. Second, the results based on derived import slopes or elasticities (estimated at 2.6 and 3.6, respectively) are close to each other and yield much lower weights than the results based on directly specified import demands. Third, the Armington model elasticities, the most widely used in previous work, yield the smallest weights on domestic welfare, although within range of the weights estimated in previous work. Fourth, based on derived import demand parameter estimates, the role of campaign contributions within the PFS model is much stronger than previously found.

The parameter estimate for the PFS model (14) with derived import demand elasticities implies a general welfare weight of approximately 2.6. This estimate is the lowest found to date via econometric estimations. In fact, it is nearly 674 times smaller than the weight obtained with direct import elasticities and 1,282 times smaller than the one using Armington elasticities.

The parameter estimate for the PFS model (13) with derived import slopes implies a general welfare weight of approximately 3.6. This parameter is 84 times smaller than the one obtained with linear import slopes using the same explanatory variables, except for the use of import prices instead of domestic prices used in the derived model. In spite of this, the direct import slope model yields weight estimates that are much lower than those using price elasticities of imports. Yet, the parameter estimates using the direct import slopes appear to be large as they indicate that the government values domestic welfare 303 times more than campaign contributions, making the latter rather irrelevant in influencing trade policy.

From the parameter estimates based on either direct import slopes or elasticities, the relative weights on general welfare are unrealistically large. These results indicate that protection is unequivocally not for sale as the weights range between 303 for direct import slopes, to 3,360 for Armington elasticities. Gawande and Bandyopadhyay (2000) found the relative welfare weight to be approximately 3,175, which is in between our elasticity-based estimates. Goldberg and Maggi (1999), Gawande and Li (2004), Eicher and Osang (2002), Mitra, Thomakos and Ulubasoglu (2002), and McCalman (2004) estimate it between 24 and 125, which are between the estimates using derived and direct import demand parameters.

Why are the derived slope or elasticity models yielding much smaller estimates for domestic welfare than the import demand models specified directly? In theory, domestic excess demand should be equivalent to imports at various post-tariff prices. In practice, given that we are using apparent consumption to measure domestic consumption, imports and excess demand are indeed equivalent but post-tariff and domestic prices are not. Domestic prices, import prices, and tariffs all come from different sources.

Note that, as shown in Table 1, domestic prices are much more stable than import prices, which is not only one of the objectives of domestic food and agricultural policies in some of the subsectors included (i.e., sugar and dairy industries) but also implies partial passthrough. While the average coefficient of variation for post-tariff prices in the industries analyzed is approximately 0.50, the one for domestic prices is approximately 0.20. In other words, the spread of import prices is 250% larger than the one for domestic prices as measured by their coefficient of variation. Furthermore, the correlation coefficient of domestic and import prices was 0.40, attesting that there is no perfect price transmission from world to domestic markets in processed food markets.

Estimating direct elasticities based on domestic rather than import prices leads to a domestic welfare weight of 71 instead of 1,766. By virtue of incorporating more information, derived estimates provide perhaps a more realistic measure of the potential rather than the actual tariff response free of noise incorporated in direct import models, particularly the Armington model which in addition assumes imperfect substitution between home and foreign food products. Finally, partial price transmission might be due to a myriad of factors not accounted for in the PFS model which may partially isolate domestic prices, such as government intervention other than import tariffs and quotas, market power, contracts, and uncertainty, among other factors.

Why are the elasticity-based empirical models for equation (14) yielding such large general welfare weights relative to the slope-based models? Part of the answer may lie on the disproportional scale of the regressors used in the PFS model. To illustrate, take equation (12) and, for simplicity, ignore buyers' lobbying. Then the level of protection is proportional to the inverse of import penetration and the inverse of the price elasticity of imports. Thus, if import penetration ratios and/or import elasticities are quite small, the regressors in the PFS model based on import elasticities (the one usually estimated) will be quite large. If this is generally the case in the sample, then the general welfare weights will tend to be quite large.

In our sample, the average import penetration ratio was less than 7% and import elasticities average somewhat over 1 (in absolute value) and are often a fraction of that, thus magnifying the proportionality problem. As seen from Table 1, the average adjusted regressor for equation (13), using import slopes, is approximately two times and 22 times larger than the average adjusted dependent variables (T_{it}^{M}) for derived and direct estimation methods, respectively. In contrast, the average adjusted regressor for equation (14), using elasticities, are 551 and 111 times larger than the adjusted dependent variables (T_{it}^{e}) using the direct and Armington import demand models.

Even though the estimated elasticities are consistent with those obtained in previous empirical work, one should keep in mind that previous estimates are viewed as too small by many trade economists (McDaniel and Balistreri, 2002). Although understandably empirical estimates are sensitive to estimation technique and misspecification (e.g., the perfectly competitive assumption), the divergence in direct or Armington estimates does not appear enough to produce welfare weights in the same range as the derived slope estimates. For instance, Gallaway, McDaniel and Rivera (2003) find that long-run Armington elasticities are about twice as large as short run ones—hardly a magnitude to overcome the proportionality problem.

The null hypothesis that the government only cares about aggregate welfare (H_0: $\beta = 0$) was rejected at the 1% level by all models. Alternative hypotheses including that the government cares equally about campaign contributions and general welfare (H_0: $\beta = 1$) and that the government only cares about campaign contributions (H_0: $\beta = 10{,}000$, using an arbitrarily large number) were also rejected at the 1% level for all model specifications. Judging from the magnitude of the welfare weights, the government mostly cares about general welfare in setting trade protection. Judging from the hypothesis tests, the government is sensitive to both aggregate welfare and campaign contributions, although obviously more so to aggregate welfare.[15]

Summary and Conclusions

This article applies the Protection for Sale model to the U.S. food processing industries using more direct measures of tariff rates and tariff equivalents and more disaggregated data than previous work, as well as alternative empirical specifications including the PFS model based on import demand slopes and the standard elasticity specification based on Armington and direct import demand models.

The empirical results strongly support the qualitative predictions of the Grossman and Helpman (1994) model with regard to the structure of trade protection. Organized sectors are granted protection and the degree of protection inversely depends on import penetration and the price elasticity of import demand. In addition, industries facing politically organized buyers are granted lower tariffs.

A surprising result is that the estimated general welfare weight is much lower than that found in previous studies when the PFS model is estimated with import slopes or elasticities derived from domestic supply and demand. This weight is found to be between 2.6 and 3.6 times the weight the government attaches to campaign contributions. However, in spite of stark differences in data set and empirical procedures, the welfare weights estimated using import slopes or elasticities based on imports--from either the Armington model or directly specified import demand models--are strikingly similar and of the same large magnitude as those of previous empirical work, implying that protection is not for sale, as the general welfare weights range between 303 and 3,606.

The results using the direct import elasticity (including Armington's) or slope specifications of the PFS model beg the question raised by Gawande and Bandyopadhyay (2000) as to why empirically the GH model yields such high weights on domestic welfare vs. campaign contributions. Our analysis suggests that the main culprit is the relatively low magnitude of the import price elasticities that result in abnormally large regressors in the PFS model. Although low import price elaticities might result from volatile import prices vs. domestic prices, the problem of dimensionality of regressors is exacerbated in the presence of low import penetration ratios that characterize most industries in our sample. This dimensionality problem can be circumvented by estimating the PFS model using import slopes or elasticities derived from domestic demand and supply using domestic prices. Whether the results of this study can be extended to industries beyond those in our sample is a question that awaits further empirical analysis.

Footnotes

[1] See Gawande and Krishna (2003) for a review of empirical work using the PFS model and other approaches. It should also be noted that there have been attempts to explain the very large weights on general welfare. For example, Gawande and Li (2004) introduce uncertainty and a low probability of obtaining the desired protection level to create lower welfare weights. As in previous attempts, however, those low welfare weights are the result of simulations and assumptions rather than econometric analysis. Protection for Sale has also been tested in countries other than the United States (Turkey by Mitra, Thomakos, and Ulubasoglu, 2002; Australia by McCalman, 2004). These studies similarly find very high weights on general welfare.

[2] The food processing industries accounted for 14% of total U.S. manufacturing value of shipments, involving 26,000 establishments and 1.5 million workers in 1992 (Connor and Schiek, 1997).

[3] Typically, the PFS model for the U.S. is tested for manufacturing industries at the 3-digit SIC level (e.g. Goldberg and Maggi, 1999; Eicher and Osang, 2002). In contrast, this article tests the PFS model using data from the U.S. food processing industries (industries in the SIC 20 classification) at the 4-digit SIC level.

[4] Equation (12) is a slight modification of the original PFS model since it does not contain the percentage of the population that is organized in lobbies. This modification allows us to focus on the welfare weight attached to general welfare, while at the same time including the influence of buyer industry lobbies that may oppose trade protection for a particular industry.

[5]Due to data availability constraints, the 1972 (instead of the 1987) SIC definitions were used. Data translation tables were used for the cases where only the 1987 SIC or USITC data were available. Although it would have been desirable to extend the analysis to more recent years, missing data on import values and especially on import prices made it impossible to include years after 1992.

[6]We used input-output tables from 1977, 1982, 1987 and 1992. The input values were then linearly interpolated for years in between for which no input-output tables are published.

[7]We are grateful to Frederick Nelson of USDA's Economic Research Service for providing updated data on tariff-rate equivalents of import quotas.

[8]We are grateful to professors Elena Lopez and Emilio Pagoulatos for furnishing their import price indices for the 1972-87 period. These price indices were extrapolated adopting their methodology (Lopez and Pagoulatos, 2002).

[9]As imports are assumed to be perfect substitutes for domestic products, import elasticities can be calculated from domestic supply and demand elasticities, given that $M_{it} = D_{it} - Q_{it}$, $e_{it} = W_{it}^{d}\eta_{it}^{d} - W_{it}^{s}\eta_{it}^{s}$, where η_{it}^{k} are the domestic price elasticities. The weights $W_{it}^{d}(= D_{it} / M_{it})$ and $W_{it}^{s}(= Q_{it} / M_{it})$ are set to their averages for each industry in order to make the import elasticities constant over time for comparison to the other import demand elasticity estimates which are also constant over time for each industry. It should be noted that making the weights variable by simply using instrumental variables for demand, supply, and imports leads to very similar results (β=0.326 vs. 0.382).

[10]For instance, the average R^2 for the home-bound domestic production, calculated from the squared correlation between the predicted and the observed values,

was 87%, ranging from 48 to 99% with a median of 90%. For imports, the average R^2 was 84%, ranging from 44 to 99% also with a median of 90%.

[11]In comparison, Goldberg and Maggi (1999) used any positive values of predicted I_i to define industries with organized sectors, using a standard (non-discrete choice) equation model. To endogenize this variable, an additional equation was specified, based on the work of Mitra (1999), Grier et al. (1991) and others. Explanatory variables include the Herfindahl index to denote industrial concentration, deflated sales to denote the size of the industry, and capital intensity (the ratio of fixed capital assets to sales).

[12] If the import slope or the elasticity estimate error is the only source of error in the equation, then one ends up with heteroskedasticity dependent only on the level of protection using (13) or (14). However, this is highly unlikely since we are using instruments for home-bound production and imports. In addition, if one assumes other sources of unknown errors to equations (11) and (12), which is highly likely, then heteroskedasticity might also depend on other unspecified factors. Therefore, we opted to use White's (1980) heteroskedasticity-consistent covariance matrix to correct for an unknown form of heteroskedasticity.

[13] Given the low percentage of zero tariff observations, the Tobit results produced parameter estimates quite close to those presented here. On a related point, Maddala (1988) advocates the use of the Tobit model when the sample is censored or truncated. In our case, zero observations correspond to actual government decisions and are, therefore, non-censored and non-truncated. Moreover, we did not find any evidence of import subsidies, which would correspond to negative tariff rates.

[14] For the industry effects, two industries (SIC 2032=canned specialties and SIC 2082=malt liquors) made the results collapse due to lack of variation of the dependent and independent variables. Thus, the observations for those industries were eliminated in the pooling tests, although their inclusion in the time effect tests did not affect the results. All F-statistics are insignificant at the 5% level for the time effects but significant for the industry effects. For instance, the individual β coefficients ranged between 0.0004 and 5.02 for the derived slope model.

[15] As observed by Gawande and Bandyopadhyay (2000), high values of a, such as the ones found in this study, imply that the relative weight placed on net aggregate welfare versus the weight placed on campaign contributions is close. If we denote W as gross aggregate welfare and W-C as net aggregate welfare, then rewrite $C + aW = (1 + a)C + a(W - C)$. As a goes to infinity, the weight on net welfare converges to the weight on campaign contributions.

References

Armington, Paul S. "A Theory of Demand for Products Distinguished by Place of Production." *IMF Staff Papers* 16(1969): 159-176.

Bartelsman, Eric J. and Wayne Gray. *The NBER Manufacturing Productivity Database.* National Bureau of Economic Research, Technical Working Paper 205, 1996.

Connor, John M. and William A. Schiek. *Food Processing: An Industrial Powerhouse in Transition.* New York: John Wiley & Son, 1997.

Eicher, Theo and Thomas Osang. "Protection for Sale: Comment." *American Economic Review* 92(2002): 1702-10.

Federal Election Commission. *FEC Reports on Financial Activity, 1977-92*, Final Reports: Vols. III and IV, Washington D.C., various years: 1977-78, 1981-82, 1987-88, and 1991-92.

Feenstra, Robert C., "NBER Trade Database, Disk1: U.S. Imports, 1972-1994: Data and Concordances." NBER Working Paper no. 5515, March 1996.

Food and Agriculture Organization of the United Nations. *FAO Statistical Databases*. Retrieved January 10, 2002 from http://apps.fao.org.

Gallaway, M.P., C.A. McDaniel, and S.A. Rivera. "Short-run and Long-run Industry Estimates of U.S. Armington Elasticities." *North American Journal of Economics and Finance* 14(2003): 49-68.

Gawande, Kishore and Usree Bandyopadhyay. "Is Protection for Sale? Evidence on the Grossman-Helpman Theory of Endogenous Protection." *The Review of Economics and Statistics* 82 (1) (2000): 139-152.

Gawande, Kishore and Pravin Krishna. "The Political Economy of Trade and Policy: Empirical Approaches." In E. Kwan Choi and James Harrigan: *Handbook of International Trade.* Oxford: Blackwell, 2003.

Gawande, Kishore and Hui Li. "The Case of the Missing Contributions." Working Paper, Bush School of Government, Texas A&M University, 2004.

Goldberg, Pinelopi K. and Giovanni Maggi. "Protection for Sale: An Empirical Investigation." *American Economic Review* 89 (1999): 1135-1155.

Grier, Kevin B., Michael C. Munger and Brian E. Roberts. "The Industrial Organization of Corporate Political Participation." *Southern Economic Journal* 57 (1991): 727-738.

Grossman, Gene M. and Elhanan Helpman. “Protection for Sale.” *American Economic Review* 84 (1994): 833-850.

Lopez, Elena and Emilio Pagoulatos. "Estimates and Determinants of Armington Elasticities for the U.S. Food Industry." *Journal of Industry, Competition and Trade* 1(2002): 247-258.

Lopez, Rigoberto A. “Campaign Contributions and Agricultural Subsidies.” *Economics of Politics* 13(2001): 257-279.

Maddala, G. A. *Introduction to Econometrics.* New York: Macmillan, 1988.

Matschke, Xenia and Shane M. Sherlund. “Do Labor Issues Matter in the Determination of U.S. Trade Policy? An Empirical Reevaluation.” Working Paper, Department of Economics, University of Connecticut, 2004.

McCalman, Phillip. “Protection for Sale and Trade Liberalization: An Empirical Investigation.” *Review of International Economics* 12(2004): 81-94.

McDaniel, Christine A. and Edward J. Balistreri. “A Discussion on Armington Trade Substitution Elasticities.” Working Paper, Office of Economics, U.S. International Trade Commisssion, Washington, D.C., 2002.

Mitra, Devashish. “Endogenous Lobby Formation and Endogenous Protection: A Long-Run Model of Trade Policy Determination.” *American Economic Review* 89 (1999): 1116-1134.

Mitra, Devashish, Dimitrios D. Thomakos, and Mehmet A. Ulubasoglu. “Protection for Sale in a Developing Country: Democracy vs. Dictatorship.” *Review of Economics and Statistics* 84(2002): 497-508.

Reinert, K.A. and R. Clinton. “Estimated Elasticities of Substitution for Analysis of a North American Free Trade Area.” U.S. International Trade Commission, Office of Economics, Staff Research Study, Washington, D.C., 1991.

Shiells, Clinton R., Robert M. Stern, and Alan V. Deardorff. “Estimates of the Elasticities of Substitution between Imports and Home Goods for the United States.” *Weltwirtschaftliches Archiv* 122 (3) (1986): 497-519.

U.S. Department of Agriculture. *Estimates of Producer and Consumer Subsidy Equivalents: Government Intervention in Agriculture, 1982-92.* Economic Research Service, Statistical Bulletin No. 913, 1994.

___________. *Foreign Agricultural Trade of the United States.* Economic Research Service, Commodity Economics Division, 1988-92.

U.S. Department of Commerce, Bureau of Economic Analysis. *The 1977, 1982, 1987, 1992 Benchmark Input-Output Accounts of the United States.* Washington, D.C., various years.

U.S. International Trade Commission. *Estimated Tariff Equivalents of U.S. Quotas on Agricultural Imports and Analysis of Competitive Conditions in U.S. and Foreign Markets for Sugar, Meat, Peanuts, Cotton, and Dairy Products.* USITC Publication No. 2276, Washington, D.C., 1990a.

___________. *The Economic Effects of Significant U.S. Import Restraints, Phase II: Agricultural Products and Natural Resources.* USITC Publication No. 2314, Washington, D.C., 1990b.

__________ *Interactive Tariff and Trade Dataweb.* Retrieved January 10, 2002 from http://dataweb.usitc.gov

White, H. "A Heteroskedasticity-Consistent Covariance Matrix Estimator and a Direct Test for Heteroskedasticity." *Econometrica* 48(1980): 149-170.

Table 1. Summary Statistics for Key Variables Used.

Variable	Notation	Mean	Std. Dev.	Min.	Max.
Specific Tariff	t_{it}	0.215	0.425	0.00	2.412
Tariff (% of Dom. Price)	τ_{it}	0.124	0.187	0.00	0.746
Organization Dummies	I_{it}	0.731	0.444	0.00	1.000
Home-Bound Prod.	$\hat{Q}_{it}$	5943.3	7571.0	126.81	38399
Org. Buyers Purchases	$\sum I_{jt} X_{it}^{j}$	698.80	1492.8	0.00	8270.9
Import Quantities	$\hat{M}_{it}$	338.43	470.78	0.809	2258.7
Import Prices	p_{it}^{M}	1.073	0.459	0.242	3.6997
Post Tariff Import Prices	$p_{it}^{M} + t_{it}$	1.288	0.658	0.266	4.072
Domestic Price	p_{it}	1.328	0.265	0.814	2.172
Derived Elasticities (a. v.)	$\hat{e}_{i,der}$	654.2	2441.3	0.295	29829
Direct Elasticities (a. v.)	$\hat{e}_{i,dir}$	1.458	1.204	0.230	4.696
Armington Elast. (a. v..)	$\hat{e}_{i,arm}$	1.265	1.114	0.190	5.102
Derived Import Slopes (a. v.)	$\hat{M}'_{i,der}$	10090	13882	212.57	45554
Direct Import Slopes (a.v.)	$\hat{M}'_{i,dir}$	556.2	1043.5	0.204	3945.8
Adj. Tariff (derived)	$T_{it,der}^{e}$	193.4	1121.0	0.00	13090
Adj. Tariff (direct)	$T_{it,dir}^{e}$	0.181	0.373	0.00	2.577
Adj. Specific Tariff (derived)	$T_{it,der}^{M}$	2249.2	8606.3	0.00	70074
Adj, Specific Tariff (direct)	$T_{it,dir}^{M}$	143.11	682.5	0.00	5413.5
Adj. Tariff (Armington)	$T_{it,arm}^{e}$	0.899	0.118	0.00	0.564
Adj. Regressor (eq. 13)	Z_{it}^{M}	4384.2	6515.7	-497.72	31764
Adj. Regressor (eq. 14)	Z_{it}^{e}	99.89	220.98	-85.501	1302.6

Note: The subscripts ‘arm’,‘dir’ and ‘der’are used to distinguish estimates based on the Armington, direct import demand models, and derived estimates based on domestic demand and supply, respectively. The term ‘a.v.’ stands for absolute value. Note that import slopes and elasticities are constant over time for each industry.

Table 2: Alternative Import Demand Estimates

SIC	Industry	Elasticities: Derived $\hat{e}_{i,der}$	Elasticities: Direct $\hat{e}_{i,dir}$	Elasticities: Armington $\hat{e}_{i,arm}$	Slopes: Derived $\hat{M}'_{i,der}$	Slopes: Direct $\hat{M}'_{i,dir}$
2011	Meat Packing Plants	-14.798	-1.554	-0.923	-43010.4	-3618.9
2013	Sausage & Prepared Meats	-5.845	-0.960	-0.733	-19681.7	-845.6
2016	Poultry Dressing Plants	-1000.8	-0.230	-0.829	-45554.2	-15.97
2017	Poultry & Egg Processing	-625.0	-2.715	-1.066	-8521.4	-11.44
2021	Creamery Butter	-1622.8	-1.679	-0.865	-4553.7	-0.20
2022	Cheese, Natural & Processed	-84.277	-0.595	-0.500	-40806.4	-107.36
2023	Condensed & Evaporated Milk	-1.359	-0.838	-0.887	-976.2	-307.01
2026	Fluid Milk	-153.48	-1.050	-0.479	-5534.1	-26.85
2032	Canned Specialties	-268.94	-2.482	-0.752	-9001.4	-81.34
2033	Canned Fruits & Vegetables	-14.361	-0.832	-1.262	-14927.5	-696.72
2034	Dried/Deh. Fruit & Veg.	-15.262	-0.305	-0.802	-1575.5	-14.24
2035	Pickled Sauces & Salad Dress.	-41.739	0.547	-2.973	-3383.5	-51.49
2046	Wet Corn Milling	-73.183	-3.403	-3.405	-3560.8	-358.85
2051	Bread & Bakery Products	-256.358	-4.696	-5.102	-23967.0	-143.47
2061	Raw Cane Sugar	-1.559	-1.384	-0.189	-212.5	-88.61
2062	Refined Sugar	-0.867	-4.066	-0.408	-825.7	-3945.8
2065	Candy & Confectionary Prod.	-8.741	-2.433	-0.882	-3947.1	-634.30
2067	Chewing Gum	-59.628	-0.760	-0.194	-767.9	-12.29
2074	Cottonseed Oil Mills	-139.232	-1.190	-1.612	-1186.3	-43.71
2076	Vegetable Oil Mills	-11.476	-0.351	-1.050	-1165.4	-1.62
2082	Malt Liquors	-0.924	-0.602	-1.519	-3522.2	-1563.04
2091	Canned & Cured Seafood	-9.267	-1.251	-1.096	-3802.9	-703.04
2095	Roasted Coffee Processors	-3.193	-0.282	-0.581	-1055.0	-57.61
2098	Macaroni & Spaghetti	-19.822	-0.803	-2.248	-622.0	-17.89
Simple Average		-654.2	-1.458	-1.265	-10090.0	-556.2

Note: The import slopes were derived from a supply and demand model while the import price elasticities were derived from double log specifications so that all the parameters are constant over time for each industry. Note that the slopes and elasticities in this table are expressed in actual instead of absolute values.

Table 3. Results From Alternative Empirical Specifications of the PFS Model, U.S. Food Manufacturing Industries, 1972-92.

Variable	Parameter	Import Elasticities: Derived	Import Elasticities: Direct	Import Elasticities: Armington	Import Slopes: Derived	Import Slopes: Direct
		$T^{e}_{it,der}$	$T^{e}_{it,dir}$	$T^{e}_{it,arm}$	$T^{M}_{it,der}$	$T^{M}_{it,dir}$
Explanatory Variable						
Z^{e}_{it}	β	0.382	0.00057	0.00030		
		(0.0681)	(0.0007)	(0.00003)		
Z^{M}_{it}	β				0.277	0.0033
					(0.069)	(0.0006)
Implied Parameters						
Relative Weight on General Welfare	a	2.621	1,766	3,360	3.606	303
Normalized Weight	$\frac{a}{1+a}$	0.724	0.9994	0.9997	0.9997	0.9967

Note: Number of observations = 360. Industries are defined at the 1972 4-digit SIC levels.
The results correspond to two-stage least squares regression with heteroskedasticity-robust standard errors.

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