Protection for Sale: An Empirical Investigation

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The Grossman-Helpman "Protection for Sale" model, concerning the political economy of trade protection, yields clear predictions for the cross-sectional structure of import barriers. Our objective is to check whether the predictions of the Grossman-Helpman model are consistent with the data and, if the model finds support, to estimate its key structural parameters. We find that the pattern of protection in the United States in 1983 is broadly consistent with the predictions of the model. A surprising finding is that the weight of welfare in the government's objective function is many times larger than the weight of contributions. (*JEL* F1)

In the last few years, trade economists have paid increasing attention to the political-economy determinants of trade policies. A prominent model in this recent literature is Gene M. Grossman and Elhanan Helpman (1994), which emphasizes the influence exerted by special-interest groups on policy makers by means of political contributions. This is a model with a relatively simple structure that yields clear-cut empirical predictions, and has been applied in a number of subsequent theoretical analyses. When a theoretical framework gains prominence and becomes a popular tool for further research, it becomes important, we believe, to check how well the model squares with the empirical evidence. The objective of the present paper is to investigate the empirical validity of the Grossman-Helpman (G-H) model and, if the model finds support, to estimate its structural parameters, such as the weight attached by the government on welfare relative to contributions.

The G-H model has strong implications for the cross-sectional structure of trade protection. In particular, it predicts that cross-sectional differences in protection should be entirely explained by three variables: import elasticity, import-penetration ratio, and whether or not the industry is politically organized. Note that, according to the model, the organized industries are the ones that contribute money to the government; thus, if contributions are observed, one can use contributions data to determine which industries are organized. The model's predictions with respect to the relevant coefficient signs are as follows: (i) trade protection should be higher in industries represented by a lobby, and in industries with a lower import elasticity; (ii) *within the subset of organized industries,* protection should be higher in industries with lower import penetration, whereas in the group of nonorganized sectors, protection should increase with import penetration.[1] We will review the theoretical model and the derivation of these results in the next section.

Strictly speaking we do not test the G-H model, because we do not have a well-specified alternative hypothesis. One possibility would be to take as an alternative the classical optimal-tariff model, in which tariffs are imposed for terms-of-trade reasons. However this model would not have much of a chance, since it is

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[1] For a comparison between these predictions and those of other political-economy models in the literature, the reader is referred to Helpman (1995). In particular, he compares Ronald Findlay and Stanislaw Wellisz (1982), Arye L. Hillman (1982), Grossman and Helpman (1994), and Wolfgang Mayer (1984) by formulating all of these models in a specific-factor framework.

grossly inconsistent with the fact that virtually all countries in the world, no matter how small, have some form of trade protection. Our objective in this paper will simply be to check whether or not the predictions of the G-H model are consistent with the data, and to estimate its structural parameters.

To probe the model, we use data on nontariff barriers for the United States in 1983.[2] The model fits the data reasonably well. The coefficient signs are the ones predicted by the model. In particular, we find that the protection pattern differs between politically organized and nonorganized sectors: within the group of nonorganized sectors, protection tends to increase with import penetration; for organized sectors, on the other hand, there is weak evidence that protection is inversely related to import penetration (the coefficient estimate on the inverse of import penetration is positive, but not statistically significant).

To further test how closely the model fits the data, we introduce more variables in the estimation, and test whether they add explanatory power to the strict G-H model. These are variables commonly used in previous studies for tariff or nontariff barrier equations (employment size, sectoral unemployment rate, measures of unionization, changes in import penetration, buyer and seller concentration, etc.). The idea is that, if any of these regressors are found to have additional explanatory power, this may be an indication that the theory provides an incomplete explanation of trade protection, and may suggest in which direction the model could be extended to improve its empirical fit. Strikingly, we find that none of the additional variables improves the explanatory power of the strict G-H model, with the possible exception of employment size and unemployment rate (they both have a nonnegligible impact on protection, but the formal likelihood ratio test does not reject the simpler specification in favor of the extended one). We also find that there is no significant difference between a version of the model that treats the political-organization dummies as endogenous and a version in which these dummies are assumed to be exogenous.

Another interesting aspect of our exercise is the estimation of the structural parameters of the model. The most noteworthy result is that the weight of welfare in the government's objective is estimated to be around 0.98 (with a 95-percent confidence interval of 0.97–0.99), as opposed to a weight of around 0.02 for contributions. This result seems consistent with the fact that trade barriers in the United States are quite low; even in 1983 the average coverage ratio was only 0.13, substantially smaller than the potential maximum of 1. One might be tempted to interpret our estimates as suggesting that the U.S. government essentially maximizes welfare, the trade policy outcome is essentially free trade, and thus the G-H model is consistent with the data but in a way that is not very interesting. However, we can reject the hypothesis that the government is a *pure* welfare maximizer, as well as the hypothesis that the model has no explanatory power. Thus we are inclined to conclude that, even though the estimated magnitude of political considerations in the government's objective is small, the G-H model has nonnegligible explanatory power for the cross-sectoral pattern of import barriers (whereas the traditional model of a welfare-maximizing government does not).

In the literature there is a large number of empirical studies that investigate the political-economy determinants of trade protection. A few examples are Edward John Ray (1981), Howard P. Marvel and Ray (1983), Robert E. Baldwin (1985), Daniel Trefler (1993), and Jong Wha Lee and Phillip Swagel (1996). We refer the reader to Dani Rodrik (1995) for a comprehensive survey of this literature. These works take a reduced-form approach, in the sense of not being guided by a theoretical model. The present paper departs from this literature in that we focus on a specific theoretical model, and let our estimation be tightly guided by it, both in the sense of including only the variables that the model suggests are relevant,

[2] We could not use data on other countries because we did not have access to data on lobby contributions or import elasticities. As for the United States, using multiple years of data is not feasible because the coverage ratios that we use as a measure of nontariff barriers are not comparable across time (there are various inconsistencies in their method of construction from year to year); comparisons across sectors for a particular year are in contrast more valid.

and of utilizing the functional forms predicted by the model.[3]

One finding that is fairly consistent across the aforementioned studies is that trade protection tends to be higher in industries with higher import penetration. This finding seems at odds with the G-H model. We argue that the discrepancy between this finding and our results is due to the different way the explanatory variables enter the estimating equation. In particular, the estimating equations employed in previous work typically introduce import-penetration and political-organization variables additively on the right-hand side, whereas we adopt the interactive specification dictated by the G-H model.

We can see two important limitations of our exercise. The first is that we use nontariff barriers, and in particular coverage ratios, whereas the model's predictions are in terms of tariff levels. The reason we use nontariff barriers is that tariffs are determined cooperatively in the GATT-WTO, and the model we are focusing on applies to situations where the government sets trade barriers noncooperatively. In principle, one could attempt to analyze the process of tariff determination by estimating Grossman and Helpman's (1995) "Trade Talks" model, in which governments set tariffs cooperatively. The problem with this approach is that an examination of the empirical implications of the model requires data on political-organization variables for the rest of the world; even if it were possible to agree on which countries the "rest of the world" should include, obtaining the necessary information on lobbying and contributions for each country would be extremely hard, if at all feasible.

There are several problems associated with the use of coverage ratios. One concern is that coverage ratios may be a very imprecise proxy for the actual restrictiveness of import barriers. However, for our qualitative cross-sectional results, we need only assume that the ranking of sectors by coverage ratios is roughly the same as their ranking by tariff equivalents, which does not seem an unreasonable assumption. The second problem concerns the role of quantitative restrictions. In particular, voluntary export restraints (VERs) are generally set in a cooperative fashion, thus it is not clear that a noncooperative model is appropriate for them. Also, the implications of quantitative restrictions can in principle be different from those of price-oriented measures, since the rents generated by them may not accrue to the government. To check that our results are not driven by the presence of quantitative restrictions in our measure of protection, we reestimated the model including only price-oriented measures (such as countervailing duties and antidumping duties), and the basic results remained unchanged. The problems associated with coverage ratios, and the way we deal with them, are discussed in more detail in Section II.

The second limitation of our exercise is that we do not have reliable data on foreign-export supply elasticities.[4] According to the model, these elasticities incorporate the purely economic determinant of tariffs, namely the terms-of-trade gains from tariffs. In ignoring these elasticities we are effectively making a small-country assumption, which may be controversial in the case of the United States. However, we believe that the assumption that a single country does not possess monopsony power is quite realistic for most markets. In the case of the United States, we feel that this assumption is reasonable with the exception of the market for oil, which is not included in our analysis.[5]

The structure of the paper is as follows. In Section I we review the G-H theoretical model. In Section II we describe the econometric specification. In Section III we describe the data, and the empirical findings. Section IV concludes.

I. Review of the Grossman-Helpman Model

In this section we briefly review Grossman and Helpman's (1994) model. We will present a

[3] Late in the revision process, we became aware of a paper by Usree Bandyopadhyay and Kishore Gawande (1998) that tests Grossman and Helpman's (1994) model. The paper reaches qualitatively similar conclusions as ours, finding that the model is broadly consistent with the data.

[4] The available estimates of import-demand elasticities are also noisy, although not as much as foreign-export supply elasticities. We explain how we deal with this issue econometrically in section II.

[5] Our data include only the manufacturing sector, for which, we believe, the assumption of no monopsony power is plausible.

slightly simpler version of the model that yields the exact same predictions. A desirable task would be to identify the *weakest* assumptions under which the same predictions hold, but we will not pursue this theoretical task here.

There is a continuum of individuals, and the population size equals one. Individuals have identical preferences, given by

$$U = c_0 + \sum_{i=1}^{n} u_i(c_i)$$

where c_i denotes consumption of good i, c_0 denotes consumption of the numeraire good, and u_i is an increasing concave function. The demand for good i implied by these preferences is denoted $d_i(p_i)$, where $d(\cdot)$ is the inverse of $u_i'(\cdot)$. The indirect utility of an individual with income y_i is given by $V_i = y_i + \sum_{i=1}^{n} s_i(p_i)$, where $s(p) = u(d(p)) - pd(p)$.

There are $n + 1$ inputs: labor and one sector-specific input for each sector. The total supply of labor has measure one. Good 0 (the numeraire) is produced one-to-one from labor, so that the wage is equal to one. Each of the other goods is produced from labor and the sector-specific input. The returns to specific factor i depend only on p_i and are denoted by $\pi_i(p_i)$. By Hotelling's lemma, $\pi_i'(p_i) = y_i(p_i)$, where $y_i(p_i)$ is the supply function of good i.

The government chooses specific trade taxes. A trade tax introduces a wedge between local price and international price: $p_i = p_i^* + t_i^s$, where t_i^s represents a specific import tariff if the good is imported, and a export subsidy if the good is exported. World prices p_i^* are exogenous. The government redistributes the revenue from trade policy in lump-sum fashion and equally to all citizens (if the revenue is negative, it is financed by lump-sum taxes).

Summing indirect utilities over all individuals, and noting that aggregate income is the sum of labor income, returns to the specific factors and tariff revenue, one obtains aggregate welfare:

$$W = 1 + \sum_{i=1}^{n} \pi_i + \sum_{i=1}^{n} t_i^s M_i + \sum_{i=1}^{n} s_i$$

where $M_i = d_i - y_i$.

Next we describe the political structure. Suppose that in some subset of sectors $L \subset \{1, 2, \cdots n\}$ the owners of specific factors are able to form a lobby. Let α_i denote the fraction of people who own specific factor i. Assume that each individual owns a unit of labor and at most one type of specific factor. Summing indirect utilities over all individuals who belong to lobby i and rearranging, we obtain lobby i's aggregate well-being:

$$W_i = \pi_i + \alpha_i\left(1 + \sum_{j=1}^{n} t_j^s M_j + \sum_{j=1}^{n} s_j\right).$$

Lobby i's objective is given by $W_i - C_i$, where C_i denotes the contributions paid to the government. The government's objective is a combination of welfare and contributions:

$$U^G = \beta W + (1 - \beta) \sum_{i \in L}^{n} C_i$$

where $\beta \in [0, 1]$ captures the weight of welfare in the government's objective.[6]

In their original formulation, Grossman and Helpman assume that the interaction between government and lobbies takes the form of a "menu auction," in the sense of B. Douglas Bernheim and Michael Whinston (1986). Here we assume a simpler mechanism that gives rise to the same trade policy outcome: a Nash bargaining game. At the Nash bargaining solution, trade policies are selected to maximize the joint surplus of all parties involved. The joint surplus is given by

$$\Omega = \beta W + (1 - \beta) \sum_{j \in L}^{n} W_j.$$

The equilibrium contributions depend in a delicate way on the specifics of the decision-making process and on the parameter values.

[6] Grossman and Helpman (1996) show that this objective function emerges in a political system in which lobbies use campaign contributions to influence the outcome of the election, and two parties compete for seats in parliament.

The model (both in the menu-auction version and in the Nash bargaining version) yields no simple prediction regarding contributions.

To find the equilibrium trade policies, one can rewrite Ω as:

$$\Omega = \beta + (1-\beta)\alpha_L + \sum_{i=1}^{n} [\beta + (1-\beta)I_i]\pi_i + \sum_{i=1}^{n} [\beta + (1-\beta)\alpha_L](t_i^s M_i + s_i)$$

where $\alpha_L \equiv \Sigma_{i \in L}\, \alpha_i$ represents the share of population that owns some specific factor, and I_i is a dummy that takes value one if $i \in L$ and zero otherwise. The first-order condition for tariff t_i^s is:

$$\frac{\partial \Omega}{\partial t_i^s} = \frac{\partial \Omega}{\partial p_i} = [\beta + (1-\beta)I_i]X_i + [\beta + (1-\beta) \times \alpha_L][-d_i + t_i^s M_i'(p_i) + M_i] = 0$$

which yields

$$(1) \qquad t_i^s = \frac{I_i - \alpha_L}{\dfrac{\beta}{1-\beta} + \alpha_L} \cdot \frac{X_i}{-M_i'}$$

where X_i represents domestic output for good i. The same formula can be expressed in terms of import elasticity and import-penetration ratio:

$$(2) \qquad \frac{t_i}{1+t_i} = \frac{I_i - \alpha_L}{\dfrac{\beta}{1-\beta} + \alpha_L} \cdot \frac{z_i}{e_i}$$

where t_i is the ad valorem tariff on good i, e_i is the import-demand elasticity of good i, and $z_i \equiv X_i/M_i$. Notice that the term $[1 - \alpha_L/(\beta/1 - \beta) + \alpha_L]$ is positive: the model predicts that, for organized sectors, the level of protection increases with X_i/M_i. The intuition for this result is that, if domestic output is larger, specific-factor owners have more to gain from an increase in the domestic price, while (for a given import-demand elasticity) the economy has less to lose from protection if the volume of imports is lower. Also, sectors characterized by higher import elasticity should receive less protection. The intuition for this is that when the import elasticity is higher, the deadweight loss from protection is higher, hence the government is less willing to grant protection. Finally notice two special cases in which the model predicts free trade. First, if the government does not care about contributions ($\beta = 1$), intuitively it has no incentive to impose trade barriers. Second, if all industries are organized ($I_i = 1$ for all i) and each citizen is represented by some lobby ($\alpha_L = 1$), then the joint surplus of all lobbies coincides with the well-being of society at large, hence free trade is the equilibrium outcome.

II. The Econometric Model

Equation (2) constitutes the basis of our empirical specification. To go from the theoretical model discussed in the previous section to the econometric model we estimate, we need to introduce an error term and specify its distribution. This error term can be thought of as a composite of variables potentially affecting protection that may have been left out of the theoretical model, and error in the measurement of the dependent variable. Since the G-H model is silent about the way the error term should enter the specification, we introduce it in a way that accommodates the estimation. Specifically, after bringing the import-demand elasticity on the left-hand side of equation (2), we enter the error term additively, so that the estimating equation becomes

$$(3) \qquad \frac{t_i}{1+t_i}e_i = \frac{I_i - \alpha_L}{\dfrac{\beta}{1-\beta} + \alpha_L}\frac{X_i}{M_i} + \epsilon_i = \gamma\frac{X_i}{M_i} + \delta I_i \frac{X_i}{M_i} + \epsilon_i$$

where $\gamma = [-\alpha_L/(\beta/1 - \beta) + \alpha_L]$, and $\delta = [1/(\beta/1 - \beta) + \alpha_L]$.

Note that this specification is only one among several alternatives. We could have, for example, introduced the error term additively in equation (2), leaving the elasticity e_i on the right-hand side. Or we could have brought in addition to the elasticity, the inverse penetration ratio (X_i/M_i) on the left-hand side, leaving only the organization dummy I_i on the right-hand side. Finally, we could have brought only (X_i/M_i) on the left-hand side, leaving the elasticity on the right-hand side. Our choice among these alternatives was guided by the following considerations.

First note that, according to the theoretical model, both the inverse penetration ratio and the elasticity (which is in general a function of price) should be thought of as endogenous variables. In addition, as we explain in the following subsections, we have good reason to believe that the elasticities are measured with error. There are in principle two ways in which we could deal with these issues; follow the general econometric practice of introducing endogenous, and measured-with-error, variables on the left-hand side [this would argue for taking both e_i and (X_i/M_i) on the left-hand side], or try to specify reduced-form equations for the above two variables and estimate the whole system using simultaneous equation techniques. The problem with this latter approach is that it is difficult to come up with a sensible reduced-form specification for elasticities—or at least a specification motivated by the theoretical model under consideration. The first approach [taking both e_i and (X_i/M_i) on the left-hand side] is feasible. We experimented with this specification in Section III, subsection C, and it did not change our results in any significant way.

It is natural in the context of our empirical model to suspect heteroskedasticity of the error term ϵ. In particular, one might suspect that the variance of the error term is related to the elasticity. We investigate this issue in the empirical section and Appendix C.

Notice the nonadditive structure of the relationship predicted by the model: the import-penetration ratio (or its inverse, to be more precise)[7] enters interactively with the political-organization dummy. Hence, the model predicts that the relationship between trade protection and import penetration depends critically on whether or not a sector is organized.

Our first task will be to estimate the parameters γ and δ and examine whether their signs are consistent with the theoretical predictions. The G-H model implies that $\gamma < 0$, $\delta > 0$, and $\gamma + \delta > 0$. We will then use our parameter estimates to compute the implied weight of welfare in the government's objective (β) and the fraction of the population represented by a lobby (α_L).

In order to estimate equation (3), we had to deal with a number of issues having to do with the four variables involved in the equation. In the following we discuss these issues in detail, focusing in sequence on the variables t_i, e_i, I_i, and (X_i/M_i).

A. *Protection Measure*

To measure protection we use coverage ratios for nontariff barriers.[8] Why do we look at nontariff barriers when the model, strictly interpreted, calls for tariffs? As mentioned in the introduction, a political-economy model with cooperative trade policy determination, like Grossman and Helpman's (1995) "Trade Talks" model, would be more appropriate in explaining the tariff structure in the United States, but data reasons prevent us from estimating such a model.[9,10]

[7] When we use the term "import penetration" in this paper, we refer to the ratio of imports to domestic output; this use of the term differs from the common definition of import penetration in the literature as the ratio of imports to domestic consumption.

[8] The original source of the nontariff barrier (NTB) data is the UNCTAD data base on trade control measures.

[9] In particular, we would need data on political organization for countries other than the United States. Such data are generally not available.

[10] One possibility, that we have not pursued, would be to use some total index of protection that combines tariffs and nontariff barriers to estimate the "Protection for Sale" model. The problem with this exercise is that it is not clear how to aggregate price measures with nonprice measures. One possibility would be to use the Trade Restrictiveness Index (TRI) discussed in James F. Anderson and J. Peter Neary (1996). Construction of this index requires, however, knowledge of the import-demand elasticities and market structure in the relevant sectors, as well as detailed information on the implementation of the trade restrictions, so that the recipients of the trade policy rents can be identified. While this information is available for specific sectors of the

Coverage ratios are a notoriously imprecise measure of nontariff barriers; however there seems to be consensus that, in the absence of reliable numbers on tariff equivalents, they are the best available measure [see Sam Laird and Alexander J. Yeats (1990) and Trefler (1993) for a detailed discussion]. A previous result by Trefler (1993) makes us optimistic about the use of coverage ratios. Trefler constructed coverage ratios for *tariffs,* and compared them to average tariffs. The correlation coefficient between the two variables was 0.78;[11] thus the cross-sectional inferences drawn from average tariffs and tariff coverage ratios are very similar. There is a small leap of faith here, but this makes us hopeful that the same is true for nontariff barriers, i.e., that the ranking of sectors in terms of coverage ratios is roughly the same as their ranking by (average) equivalent tariffs. If this is the case, then our qualitative cross-sectoral results should not be distorted by the use of coverage ratios.

Our more quantitative findings (in particular, the point estimates of the structural parameters) are certainly affected by the use of coverage ratios. The coverage ratio for industry i is defined as $\sum_k n_k^i w_k^i$, where the summation is taken over all the products in industry i,[12] the indicator variable n_k^i takes value one if product k is covered by some nontariff barrier, and the weight w_k^i is the import share of product k relative to total imports in the industry. Both components of coverage ratios, the weights and the zero-one measures of protection, are problematic. In the following we discuss the problems associated with each component in detail.

Using import shares as weights has the well-known shortcoming of potentially attaching low weight to products that are highly protected, since these sectors are likely to have low imports. However we should point out that this procedure is exactly the same as the one used in the computation of average tariffs; unless one is willing to work at the extremely disaggregate level tariffs are specified at, there is no obvious way to improve on this dimension. Lee and Swagel (1996) provide some encouraging results pertaining to this issue. They constructed coverage ratios for nontariff barriers using two alternative sets of weights: import shares and production shares. Just as the use of import shares tends to understate protection, the use of production shares should overstate it. The authors report that the correlation between the protection measures obtained by the two different weighting schemes is in the order of 0.98.

The second source of imprecision is linked to the use of a zero-one measure of protection for each product. One way to think about this problem is that there are two implicit assumptions in using coverage ratios as the dependent variable. The first one is that the coverage ratio is proportional to the underlying (average) equivalent tariff. This is a strong assumption, and we have no way of relaxing it. The second assumption concerns the exact factor of proportionality. One option is to assume that for coverage ratios lower than 1, the coverage ratio reflects the equivalent tariff (meaning, for example, that a coverage ratio of 0.5 corresponds to an equivalent tariff of 0.5), and equivalent tariffs higher than 100 percent are mapped onto a coverage ratio of 1. However, the few numbers that are available on equivalent tariffs suggest that this mapping may be understating the extent of protection. As an example consider the automobile industry. Goldberg (1995) estimated the tariff equivalent of the VERs on Japanese imports to be around 60 percent. In those data, the coverage ratio for autos in 1983 was 7 percent; this suggests that here the coverage ratio largely understates protection. To address this issue, we

U.S. economy (Anderson and Neary have, for example, successfully used such information to compute the TRI for quotas on Hong Kong textile exports to the United States), it is typically not available at the scale and level of aggregation needed for the present study. At any rate, we note that tariffs in the United States are very low (the average tariff is about 5 percent) and vary little across sectors, whereas nontariff barriers are higher (the average coverage ratio in our data is 13 percent) and vary considerably across sectors. In addition, we suspect that coverage ratios understate the actual extent of protection (see the discussion in this section); thus the discrepancy between the magnitude of tariff and nontariff protection may be even larger. For these reasons, we feel that accounting for tariffs would not substantially change our qualitative results. An experiment that we did perform was to include tariffs on the right-hand side. One would expect that, if the tariff component of protection were important, introducing tariffs as explanatory variables would have a strong effect. But our results were hardly affected by this.

[11] See Trefler (1993 p. 156).

[12] The product here is defined at the most disaggregated level possible, i.e., at the tariff-line level. There are around 10,000 products at this level.

allow the equivalent tariff to be a multiple of the coverage ratio (for coverage ratios less than one). In the model estimation, we experiment with different scaling factors, ranging from 1 to 3. As we will show later, allowing for this scaling factor hardly affects the qualitative results; all parameter estimates preserve their signs, and the implied share of the population represented by a lobby remains essentially unaltered. It does, however, affect the implied weight on national welfare; naturally, the higher the scaling factor, the higher our measure of protection, and the lower the implied weight the government attaches to the country's welfare.

Two other possible problems concern the presence of quantitative restrictions among nontariff barriers: First, the cross-sectional predictions may differ according to whether tariffs or quantitative restrictions are used. This is confirmed by Maggi and Andres Rodriguez-Clare (1999), who show that the predictions of the G-H model *may* change if quantitative instruments are used and if parameters fall into a certain region. Second, VERs are generally negotiated between the United States and foreign countries, thus a noncooperative model might not be appropriate for them. To deal with these concerns, we reestimated the model including only price-oriented measures (such as antidumping duties and countervailing duties) in the dependent variable, and we obtained the same qualitative results, as we mention in Section III, subsection C.

Finally, the use of coverage ratios instead of tariffs has implications for the empirical implementation of the model. Given that coverage ratios can only take values between 0 and 1, the protection variable t_i is censored. Accounting for the censoring of the coverage ratios at zero is particularly important, as 39 percent of our observations are at zero (in contrast, only one observation is associated with a coverage ratio of 1). To this end, the protection equation is specified as a Tobit with double censoring.

B. *Elasticities*

Existing estimates of trade elasticities are generally considered unreliable; apart from changing considerably from study to study, they sometimes have the wrong sign, and their standard errors tend to be large. We considered three alternative approaches of dealing with this problem. The first one was to estimate the elasticities ourselves. After reading the existing literature, however, we realized that there was no obvious way to improve on existing estimates. The majority of previous studies have employed sophisticated econometric techniques; the imprecision in the estimates is mainly the result of noisy data, with the noise increasing in the level of disaggregation. Given that we are interested in carrying out the analysis at the most disaggregate level possible, we abandoned this approach. The second alternative was to employ available estimates of trade elasticities on the right-hand side of equation (2), but correct the standard errors of our estimates to take into account the fact that the numbers were estimated. One problem with this alternative is that we do not *exactly* know how the elasticity estimates were obtained in previous studies. More importantly, this would not address the issue that elasticities should be treated as econometrically endogenous; as noted above, specifying a reduced-form equation for elasticities is rather unappealing, as we have no guidance at all what variables to include in such an equation. For these reasons, we adopted a third alternative: utilize existing trade elasticity estimates, but introduce them on the left-hand side of the estimating equation.

C. *Political-Organization Dummies*

To construct the political-organization dummies, I_i, we use data on political action committee (PAC) campaign contributions. According to the strict version of the model, one should be able to infer the set of organized industries simply by looking at contribution levels: if the contribution level is positive the industry should be organized. In our data, contribution levels are positive for all 3-digit SIC industries, so that a literal interpretation of the model would imply that all sectors in the economy are organized. However, this implication would be valid only if contributions were made exclusively to influence trade policies, and if they were measured without noise. In reality, firms contribute for a variety of other reasons, in particular to influence domestic policy. Some industries may hire lobbyists to influence domestic policies, but not make an organized effort on the trade front. Moreover there are imprecisions related to the assignment of contributions to 3-digit industries.[13] The presence of contributions extraneous to trade policy calls for a more flexible criterion in assigning the political-organization dummies. We adopted the following intuitive method: if the contribution level was below a certain threshold, the political-organization dummy was set to zero; if contributions exceeded the threshold, the dummy was set to one. We first adopted a threshold level suggested by a natural break in the data, then we experimented with different levels of the threshold. In Section III, subsection B, we will be more specific regarding these experiments.[14]

In the G-H model, the organization dummies I_i are exogenous. However, since we use contribution data to assign the political-organization dummies, and contributions are endogenous in the G-H framework, in the empirical implementation of the model we treat I_i as *econometrically* endogenous, and specify a reduced-form equation for it. In the right-hand side of this reduced-form equation we include a set of traditional political-economy regressors (concentration indices, minimum efficient scale, unionization, geographic concentration, etc.) which are natural instruments for contributions and organization dummies, as well as the exogenous variables that enter the import-penetration equation (see next subsection). To examine whether this treatment of I_i had an impact on the results, we also estimated a specification in which the political-organization dummies were treated as econometrically exogenous; this turned out to make no appreciable difference, either for the point estimates or the standard errors.

[13] The original contributions data are at the firm level, and the assignment of firms to 3-digit industries is subject to imprecisions due to the presence of multiproduct firms. The reader is referred to the appendix of Gawande's (1995) paper for a detailed description of the way contributions are assigned to industries.

[14] An assumption that is implicit in the empirical implementation of the G-H model is that the relevant unit for political organization is a 3-digit SIC industry. In other words, we are ruling out that, within the same 3-digit industry, some subindustries are organized and other are not.

D. *Import Penetration*

Both theory and previous empirical results (Ray 1981; Trefler 1993) suggest that the import-penetration ratio should be treated as endogenous. In the G-H model, trade flows, and hence import penetration, are determined just as in the specific-factors model. Even though "testing" the part of the G-H model that concerns the determination of trade flows is not the focus of our analysis, we believe that it is important to specify a reduced-form equation for the inverse penetration ratio that is consistent with the spirit of the model. To this end, we employ a specification similar to Trefler's (1993) import equation, where the import-penetration ratio is a function of factor shares in each sector. These factor shares are essentially measures of the amounts of capital, land, and skilled labor used in each sector. In addition, we include variables that may affect the propensity of a sector to get organized on the right-hand side (concentration indices, minimum efficient scale, unionization, geographic concentration, etc.). The reason these variables are included in this reduced-form specification is that import penetration depends on the level of protection which, according to the theory, is affected by political organization.

E. *The Full Econometric Model*

To formalize the discussion above, the empirical model we estimate has the following form:

$$y_i^* = \frac{t_i^* e_i}{1 + t_i^*} = \gamma \frac{X_i}{M_i} + \delta I_i \frac{X_i}{M_i} + \epsilon_i \tag{4}$$

$$t_i = \begin{cases} \dfrac{1}{\mu} t_i^* & \text{if } 0 < t_i^* < \mu \\ 0 & \text{if } t_i^* \leq 0 \\ 1 & \text{if } t_i^* \geq \mu \end{cases} \tag{5}$$

$$\frac{X_i}{M_i} = \zeta_1' \mathbf{Z}_{1i} + u_{1i} \tag{6}$$

$$I_i^* = \zeta_2' \mathbf{Z}_{2i} + u_{2i} \tag{7}$$

$$I_i = \begin{cases} 1 & \text{if } I_i^* > 0 \\ 0 & \text{if } I_i^* \leq 0 \end{cases}. \tag{8}$$

The variable t_i^* is a latent variable that can be thought of as the "true" level of protection; this is equal to a multiple (μ) of the coverage ratio if the coverage ratio is strictly between zero and one. It is important to note that we do not attempt to estimate μ as a parameter in the model. We believe this would be a questionable approach, not only because μ would appear on the left-hand side, so that the likelihood function could not be defined in the conventional way, but also because there is nothing really in the data to identify μ—it would be strictly identified off the nonlinearity in the model. Rather, we use μ to conduct a kind of sensitivity analysis regarding the magnitude of the structural parameters as a function of the restrictiveness of trade barriers. In the implementation of the model, we experiment with μ values of 1, 2, and 3; these give us a good sense of the direction in which the results change when protection increases. Similarly, I_i^* is a latent variable; if this is positive, the sector is organized, and the organization dummy takes the value 1; otherwise the dummy is zero. The vectors $\mathbf{Z}_{1i}$ and $\mathbf{Z}_{2i}$ consist of variables employed in the specifications for the inverse import-penetration ratio and the political organization dummy, respectively (in the specifications we report the two vectors $\mathbf{Z}_{1i}$ and $\mathbf{Z}_{2i}$ are actually identical). The error terms ϵ, u_1, and u_2 are assumed to be distributed as $N \sim (0, \Sigma)$.

The system (4)–(8) is estimated by maximum likelihood. To test whether the model predictions are borne by the data, we use three criteria. First, we examine whether the signs of the coefficients γ and δ are the ones predicted by theory. Second, we derive the structural parameters β and α_L for various values of the scaling factor μ, and check whether their values fall in the admissible range, i.e., between 0 and 1. Third, we successively introduce more variables in the estimation, and test whether these variables have additional explanatory power, in the sense of providing a better fit. These variables consist of the common regressors in tariff or nontariff barrier equations of previous studies (sectoral employment size and unemployment rate, unionization measures, changes in import penetration, market concentration indices, etc.). In doing this, we depart from the framework of the model, as there is no theoretical foundation for including such variables in the estimation. To the extent that any of these regressors are found to have additional explanatory power, we may get an idea of which aspects of trade protection are incompletely explained by the theory, and in which direction the model should be extended. The results are reported below.

III. Empirical Results

A. *Data*

The static nature of the model calls for a cross-sectional analysis. The data we use in our analysis refer to 1983, and are aggregated up to the 3-digit SIC level. This aggregation level was chosen to match the most disaggregate elasticity estimates available in the literature.

Import-demand elasticities are taken from Clinton R. Shiells et al. (1986), the only study that estimated elasticities at the 3-digit SIC level. While the estimates present many of the problems discussed above, they were obtained with a sound econometric methodology, and are considered among the most reliable ones in the trade literature.

The inverse import-penetration ratio (X_i/M_i) is measured as the ratio of value of shipments over imports in each industry. Both series are taken from the National Bureau of Economic Research Trade and Immigration data file.

Contributions data for the 1981–1982 and 1983–1984 congressional elections were kindly provided by Gawande. Given that there is little variation in the contribution levels across the two periods, we only use the 1981–1982 data in the estimation.

The variables employed in the inverse penetration ratio and political-organization equations were kindly provided by Trefler, and we refer the reader to his 1993 publication for a detailed discussion of them [Trefler (1993), Table 1 p. 140, and Data Appendix]. The factor-share regressors include shares for physical capital, inventories, engineers and scientists, white-collar workers, skilled labor, semiskilled labor, land (cropland, pasture, and forest), and subsoil (coal, petroleum, minerals). The variables related to political organization include seller concentration, buyer concentration, seller number of firms, buyer number of firms, minimum efficient scale, capital stock, geographic concentration, unionization, industry unemployment rate[15] and employment size, tenure, and industry growth.

B. *Results*

Since the results from the import-penetration and political-organization equations are not the main focus of the paper, we report them in Appendix A (Tables A1 and A2). The results are generally sensible, and consistent with the ones of previous, reduced-form studies. In Table A1 (import-penetration equation), the positive signs of physical capital and white-collar labor indicate that capital and/or human-capital intensive industries tend to have lower import-penetration ratios; this is consistent with the view that high-tech sectors in the United States are competitive in international markets. Most of the coefficients referring to land and subsoil shares are insignificant; this is also intuitive, as our estimation focuses on manufacturing, and land should be irrelevant for imports in manufacturing. The plausibility of the results of the political-organization equation is harder to judge, given that existing theories of political organization do not yield unambiguous predictions regarding the signs of the relevant coefficients—rather, they merely indicate which variables could affect political organization, and should therefore be included in the estimation; according to our results the main determinants of political organization include geographic concentration, the proportion of skilled and semiskilled labor, tenure, and variables that proxy for entry barriers such as minimum efficient scale and capital stock.

The results from estimating the trade protection equation are reported in Table 1. We start by estimating the system (4)–(8). To test for heteroskedasticity in the residual ϵ, we employed a conditional moment test similar to the one discussed in Andrew Chesher and Margaret Irish (1987). The test is described in detail in Appendix C. The basic idea is to test for correlations between the square of the generalized residual in (4) and a predetermined set of explanatory variables. Our previous discussion suggests that the variance of ϵ may be related to the elasticity, or the precision of the elasticity, estimates we borrowed from the literature. Accordingly, two obvious variables to include in this predetermined set are the elasticity estimates, and the standard errors of these estimates as reported in Shiells et al. (1986). If only e_i is included in the set, the test yields a $\chi^2(1)$ statistic of 1.41, thus failing to reject the null of homoskedasticity; if both e_i and the standard errors of the elasticity estimates are included, the $\chi^2(2)$ statistic is 3.72, a number again too small to reject homoskedasticity. These results are particularly reassuring as recent work suggests that these types of test tend to reject in small samples, even if the null is correct.

The results we report in this section were derived using a threshold level of \$100,000,000 in 3-digit-industry contributions to assign the political-organization dummy. This threshold was chosen because there seems to be a natural break in the data around that point; in particular, there are many sectors contributing \$130,000,000 and higher, and many sectors contributing \$90,000,000 or less, but very few between 90 and 130 million. This break appears clearly in the bar chart of PAC contributions [Figure B1] in Appendix B. In the same Appendix, we provide a list of all 3-digit SIC industries that are considered to be unorganized according to our criterion [Table B1]; the industries are sorted by the magnitude of their contributions, starting with the sectors with the lowest contributions. As evident from this list, our classification is generally consistent with common wisdom. The industries with the

TABLE 1—RESULTS FROM THE BASIC SPECIFICATION (G-H MODEL)

Variable	$\mu = 1$	$\mu = 2$	$\mu = 3$
X_i/M_i	−0.0093	−0.0133	−0.0155
	(0.0040)	(0.0059)	(0.0070)
$(X_i/M_i) * I_i$	0.0106	0.0155	0.0186
	(0.0053)	(0.0077)	(0.0093)
Implied β	0.986	0.984	0.981
	(0.005)	(0.007)	(0.009)
Implied α_L	0.883	0.858	0.840
	(0.223)	(0.217)	(0.214)

[15] The sectoral unemployment rate is based on data from the March 1983 Current Population Survey (CPS); a worker is considered unemployed in a particular industry if his/her longest job between March 1982 and March 1983 was in that industry.

lowest contributions include various subcategories of leather products, musical instruments, toys, and publishing and printing; in contrast, the industries with the largest contributions (not shown in the list), that are hence considered to be organized, refer mainly to machinery, chemicals and allied products, and transportation equipment. In the next section we examine the robustness of the results to alternative ways of constructing I_i.

Table 1 reports results for three cases: $\mu = 1$, $\mu = 2$, and $\mu = 3$. In all three cases, the signs and the t-statistics of the coefficients γ and δ are consistent with the predictions of the G-H model. The third sign prediction, $\gamma + \delta > 0$, finds weak support: the estimate of $\gamma + \delta$ is positive, but not statistically significant. Hence, our findings support the model's prediction that the relationship between protection and import penetration depends on whether or not the sector is politically organized; the positive sign and the statistical significance of the parameter δ indicate that there is a distinct pattern of protection in organized versus nonorganized sectors. The results also support the prediction that the relationship between import penetration and protection is positive within the set of nonorganized sectors. The third prediction of the model, that the above relationship is negative within the set of organized sectors, finds only weak support.[16]

These findings invite a comparison with the results of previous empirical work that typically finds a positive correlation between protection and import penetration (e.g., Edward E. Leamer, 1988; Trefler, 1993; Lee and Swagel, 1996). Is there a discrepancy between these results and ours? If so, what are its sources?

We believe that our results are compatible with the ones obtained earlier. When comparing results, however, it is important to note that the estimating equations employed in previous work typically introduce import-penetration and political-organization variables *additively* on the right-hand side. In the following we argue intuitively that, if the model is true but one uses the additive (mis)specification, one may well find a negative coefficient for X_i/M_i. If the model is true, in the subset of organized sectors we have $(t_i^*e_i/1 + t_i^*) = (\gamma + \delta)(X_i/M_i)$, with $\gamma + \delta > 0$, and in the subset of unorganized sectors we have $(t_i^*e_i/1 + t_i^*) = \gamma(X_i/M_i)$, with $\gamma < 0$. Now suppose one imposes the additive (mis)specification $(t_i^*e_i/1 + t_i^*) = \phi(X_i/M_i) + \zeta I_i + u_i$. This amounts to imposing the undue restriction that the coefficient of (X_i/M_i) be the same across the two subsets of sectors (organized and unorganized), which implies that the expected estimate of ϕ is some average of γ and $(\gamma + \delta)$. Since γ and $(\gamma + \delta)$ have opposite signs, for certain configurations of data this average can be negative, while if one utilizes the correct interactive specification one would find a positive coefficient $(\gamma + \delta)$.[17]

Thus, the positive correlation between import penetration and protection documented in previous papers does not constitute evidence against the model. To examine the validity of this interpretation, we estimated the protection equation without interacting import penetration with political organization, and did indeed replicate the well-documented positive relationship between imports and protection. In addition, we estimated a specification in which the right-hand side includes a constant, the two additive terms, and the interaction term. The results from this specification are reported in Table A3. Note that while the two coefficients dictated by the theoretical model are significant and have the expected signs, the constant and the coefficient on I_i are insignificant (even though the latter coefficient has the expected sign indicating that organized sectors are likely to receive higher protection). These estimates further support our interpretation of the results in the basic specification.

The parameter estimates for γ and δ can be

[16] There might appear to be an inconsistency between the data and the G-H model, in that the model predicts negative protection for unorganized sectors $[(t_ie_i/1 + t_i) = (-\alpha_L/(\beta/1 - \beta) + \alpha_L) \cdot (X_i/M_i)$ if $I_i = 0]$, while we do not observe negative protection. But it should be kept in mind that coverage ratios are positive by definition; whether or not there is some form of negative protection for some sectors, the data cannot tell us. On the other hand, the fact that we have a number of unorganized sectors with strictly positive coverage ratios is consistent with the stochastic version of the model (e.g., the presence of an additive error term on the right-hand side).

[17] Incidentally, notice that according to our estimates of γ and δ, the arithmetic mean of γ and $(\gamma + \delta)$ is a negative number.

used to retrieve the structural parameters β and α_L.[18] Naturally, the parameter values depend on the mapping of coverage ratios to equivalent tariffs, reflected in the scaling factor μ. According to the results in Table 1, if μ is set to 1, the weight of welfare in the government's objective (β) is 0.986 while the fraction of the population represented by a lobby (α_L) is 0.883; if μ is 2, the above estimates change to 0.984 and 0.858 respectively; for $\mu = 3$, they become 0.981 and 0.840, and so on. The way these estimates change as a function of μ is quite intuitive: the higher μ, the higher the equivalent tariff; this, in turn, implies a lower weight on welfare, and a lower degree of lobby representation.[19] Note that, for all these values of μ, the estimates of the structural parameters lie within the admissible range (both β and α_L are between 0 and 1), even though we did not impose any restrictions on the empirical specification to guarantee this result. Note also that in all cases both β and α_L are significantly different from zero.

An interesting feature of our results is that, independently of the value of μ, the parameter estimate for β is always very high (around 0.98), with a 95-percent confidence interval of (0.97, 0.99). The point estimate for α_L also appears to be high, but the 95-percent confidence interval is here wider; in the extreme case of $\mu = 3$, for example, this confidence interval covers values between 0.41 and 1.26. Thus it seems safe to conclude that, while our results do not allow very precise statements about α_L, they strongly suggest that welfare considerations, as captured by the parameter β, figure prominently in the government's objective.

Our interpretation of this finding is that the United States is relatively open to trade, even when nontariff barriers are accounted for. The average coverage ratio in our data, for example, is 0.13, much lower than the potential maximum of 1. The observed low protection levels can be explained within the framework of the G-H model only if welfare carries a strong weight in the government's payoff. Given that the estimated importance of political considerations in the government's objective is small, one might wonder whether we can reject the hypothesis that the government is a pure welfare maximizer ($\beta = 1$). The answer is yes; even the 99-percent confidence interval does not include $\beta = 1$.

Our results so far are broadly consistent with the theory, in the sense that the variables included in the G-H model appear to affect protection the way the model predicts they should. Next, we focus on the theoretical implications concerning variables that *should not* influence protection; the strict interpretation of the model implies that once political-organization, import-penetration, and trade elasticities are accounted for, no other observables should help explain protection. To examine this implication, we extend the empirical specification to include variables that may affect protection, but were left out of the model, and test the hypothesis that these variables have additional explanatory power. In the ideal case, these empirical extensions would be guided by well-specified alternative hypotheses, suggesting the list of regressors and the functional forms. In the absence of such alternative theories, we rely on findings of earlier studies or simple economic intuition to inform our specifications. Since the numerical values of the structural parameters β and α_L do not have any meaning in this atheoretical framework, we concentrate our discussion on the case $\mu = 1$.

A subset of the results from this exercise are reported in Table 2. The first column reports results from a specification in which both γ and δ are left out; only a constant is included. The second column replicates the results reported earlier from estimating the strict version of the G-H model, to provide a standard for comparison. The third column reports a specification that includes a constant, and the two variables suggested by the model. A comparison between columns 1 and 3 gives us some idea about the fit of the G-H model. The results indicate that the two variables indicated by the model indeed

[18] Straightforward algebra shows that $\alpha_L = -(\gamma/\delta)$ and $\beta = (1 + \gamma/1 + \gamma + \delta)$. Note that α_L and β are identified unless $\delta = \gamma = 0$ (in which case α_L is indeterminate), or $\delta = 0$ and $\gamma = -1$ (in which case β is indeterminate). But as we saw, δ is significantly different from zero, therefore no identification problems arise.

[19] Note, however, that these changes in the parameter values as a function of μ are not very large; the parameter estimates for β only change at the third digit, while the estimates for α_L change at the second digit.

TABLE 2—ALTERNATIVE SPECIFICATIONS ($\mu = 1$)

Variable	Specification 1 Log-likelihood: −134.9	Specification 2 Log-likelihood: −132.06	Specification 3 Log-likelihood: −132.04	Specification 4 Log-likelihood: −130.61
X_i/M_i	—	−0.0093	−0.0096	−0.0109
		(0.0040)	(0.0043)	(0.0045)
$(X_i/M_i) * I_i$	—	0.0106	0.0105	0.0123
		(0.0053)	(0.0053)	(0.0055)
Constant	−0.0640	—	−0.0287	−0.2619
	(0.1104)		(0.1375)	(0.2559)
Unemployment	—	—	—	1.5722
				(1.5884)
Employment size	—	—	—	1.1836
				(0.8235)

Note: Dependent variable: $(t_i^* e_i / 1 + t_i^*)$.

significantly improve the fit of the model; the log-likelihood rises from −134.9 to −132.04 once these two variables are included.

Next, we experimented with various specifications in which we successively introduced more regressors, such as changes in import penetration, various measures of employment conditions in the industry, growth, and concentration indices. The striking feature of the results was that none of the alternative specifications was found to significantly improve the fit of the model. Pairwise likelihood ratio tests failed to reject the hypothesis that the restricted version of our empirical model—the one corresponding to the strict interpretation of the G-H model—was the right one. The only specification that improved the likelihood function by a nonnegligible amount was the one reported in column 4, in which sectoral unemployment and employment size were included in the estimation. Even though the formal likelihood ratio test does not reject Specification 2 in favor of 4, the log-likelihood function improves by almost three points, and the three additional regressors are "almost significant." Note that the signs of the additional coefficients are intuitive, as one would generally expect sectors providing many jobs to receive more protection, especially when threatened by unemployment. Hence, there is some evidence that factors linked to unemployment may affect protection through channels different than the ones suggested by the G-H theory. Presumably, industries that are experiencing high unemployment rates tend to be more vocal in the political arena, and manage to obtain more trade protection. This suggests that it might be fruitful, from an empirical standpoint, to extend the G-H model to allow for sectoral unemployment and examine the impact that this has on the lobbying process.

Before concluding the section, we wish to address a possible criticism to our empirical strategy. We "test" the G-H model's prediction that the variables X_i/M_i and I_i are sufficient to explain the variation in protection, by comparing the basic specification that includes only these two variables with alternative specifications that include other variables traditionally employed in reduced-form studies. One could object that our basic specification includes not only X_i/M_i and I_i, but also all the traditional regressors, indirectly through our reduced-form equations for import penetration and political organization. There are two responses to this criticism. First, we employ the traditional regressors only to deal with an econometric endogeneity problem (in particular, the I_i dummy is endogenous because we construct it using data on contributions, which are endogenous in the model). In fact, our results are very similar if we treat I_i as econometrically exogenous, in which case we do not estimate the probit equation (7), and political-organization determinants do not appear in the estimation of the G-H model. Second, the G-H model is perfectly compatible with I_i being determined by exogenous variables that are not included in the model. If I_i is determined by a vector of exogenous variables $\mathbf{Z}$, the model predicts that I_i is a sufficient statistic for all the variables in $\mathbf{Z}$, for the purposes of explaining protection—this is one of the predictions we test.

TABLE 3—RESULTS FROM USING ALTERNATIVE THRESHOLDS TO DEFINE THE POLITICAL-ORGANIZATION DUMMY

	Thresholds		
Variable	\$50,000,000 Percent of organized sectors: 74	0.1 percent of total contributions Percent of organized sectors: 84	0.1 percent of value added Percent of organized sectors: 85
X_i/M_i	−0.0090	−0.1475	−0.0045
	(0.0039)	(0.0664)	(0.0025)
$(X_i/M_i) * I_i$	0.0099	0.1286	0.0075
	(0.0054)	(0.0697)	(0.0074)

C. *Sensitivity Analysis*

In this section we explore the robustness of our findings in four ways. First, we examine whether our results are affected by the presence of elasticities on the left-hand side. Second, we consider alternative definitions of the political dummy. Third, we compare the results from our specification, in which the political dummies are treated as endogenous, to a specification that considers them econometrically exogenous. Finally, we explore some alternative ways of defining the dependent variable.

As discussed in Section II, the measurement problems associated with trade elasticities should not bias our results given that elasticities do not appear on the right-hand side of the estimating equation. Nevertheless, to be certain that our findings were not driven by imprecisely measured elasticities, we also estimated equation (4) omitting elasticities from the specification. The signs of the parameters γ and δ remained unchanged, even though their values naturally changed. In addition, we considered two alternative ways of dealing with noisy elasticity values. The first one was to confine estimation to the sectors for which elasticity estimates had plausible sign; this meant eliminating 11, out of 107, sectors from the estimation. The second approach was to replace elasticity estimates that had the wrong sign (i.e., positive import-demand elasticities) with negative numbers very close to zero (−0.000001). Both approaches produced estimates that were favorable to the G-H model, and significantly more precisely estimated than the ones reported earlier.

Next, we examined the robustness of our results to different ways of defining the contribution threshold for the construction of the I_i dummies, both in terms of *units* in which we define the threshold, and of the critical *level*. We tried two alternative measure units, besides the absolute value of each sector's contributions: the share of the sector's contributions in total contributions and the ratio of the sector's contribution to its value added. The alternative threshold levels we experimented with were the following: contribution values of 10 and 50 million dollars, shares of 0.001, 0.005, and 0.01 of the sector in total contributions, and ratios of 0.001, 0.005, and 0.01 of the sector's contribution to its value added. The lower the threshold, the higher the fraction of sectors considered politically organized. While the parameter estimates had in all specifications the expected signs, the precision with which they were estimated declined as the fraction of organized sectors approached the limit of 1. Intuitively, to identify the parameters of interest γ and δ, we need enough variation in the political-organization dummy, for the interaction term $(X_i/M_i) * I_i$ to represent a different regressor from (X_i/M_i). Using very low or very high thresholds for constructing the dummy wipes out the variation in the independent variable, increasing the standard errors of our estimates. We report a subset of the results from our experimentation with different thresholds in Table 3.[20]

We also considered an alternative specification in which contribution levels are directly interacted

[20] It has been suggested to us that the absolute contribution levels are positively correlated with industry size; thus, political-organization dummies constructed on the basis of absolute contribution levels are going to be correlated with industry size. In the language of the G-H model, this means that larger industries are more likely to be organized. But note that we already control for size through X_i. It has also been suggested to us that, since the term $(X_i/M_i) * I_i$ is correlated with (X_i^2/M_i), our estimates of γ and δ are consistent with an alternative explanation in which protection increases with import penetration, but at a decreasing rate. To investigate this interpretation, we also estimated a specification in which in addition to (X_i/M_i) and $(X_i/M_i) * I_i$, we included the square of (X_i/M_i) on the right-hand side. The coefficient on the square of the import penetration was very small and statistically insignificant (0.00002 with a t-statistic of 0.662), while the first two coefficients and their standard errors were approximately the same as before.

TABLE 4—RESULTS FROM INTERACTING CONTRIBUTION LEVELS WITH IMPORT PENETRATION

Variable	Coefficient (Standard error)
X_i/M_i	−0.0096 (0.0036)
$(X_i/M_i) * C_i$	0.0032 (0.0013)

with the inverse import-penetration ratio. This specification departs from the theoretical framework, so that the parameters γ and δ have no structural interpretation any more, but it provides a further way of checking whether our results are sensitive to alternative definitions of political organization. To estimate this specification, equations (7) and (8) in the econometric model of Section II, were replaced by a continuous equation with contribution levels as the dependent variable,

$$C_i = \zeta'_{2c}\mathbf{Z}_{2i} + u_{2i}$$

and the estimating equation became

$$\frac{t_i^* e_i}{1 + t_i^*} = \gamma' \frac{X_i}{M_i} + \delta' C_i \frac{X_i}{M_i} + \epsilon_i.$$

The results from estimating the modified model, reported in Table 4, are consistent with the ones obtained earlier, indicating that the signs of the parameters γ and δ are not sensitive to the particular way we construct the dummy I_i.

Next, we compared the results we obtained by treating the political-organization dummies as endogenous [i.e., system (4)–(8)], to the results from a specification that treats the political dummies as exogenous [i.e., system (4)–(6)]. There was hardly a difference; this suggests that thinking of the set of organized sectors as exogenous in the short run, as the G-H model does, may be a reasonable abstraction.

We also estimated the model allowing for two different ways of defining the dependent variable. First, instead of putting $(t_i^*/1 + t_i^*)$ on the left-hand side (thus assuming that the coverage ratio is proportional to the equivalent tariff), we replaced this whole expression by t_i^*. This offers two advantages. First, it makes our specification more comparable to earlier studies, as we now use the same dependent variable [t_i^* instead of $(t_i^*/1 + t_i^*)$]. Second, it considerably simplifies the implementation of the heteroskedasticity test (see Appendix C). The results from this specification and their implications for the structural parameters are very similar to the ones we obtain by estimating the original specification [γ: −0.0147 (0.0053); δ: 0.1322 (0.0630)]. This is to be expected; because t_i is very small in our data, the denominator in $(t_i/1 + t_i)$ is close to 1 so that t_i provides a quite good approximation of $(t_i/1 + t_i)$. Our second experiment involved bringing (X_i/M_i) on the left-hand side along the lines discussed earlier in Section II, thus defining the dependent variable as $(t_i/1 + t_i)e_i(M_i/X_i)$.[21] Again, the parameter signs and their statistical significance are very similar to earlier specifications [γ: −0.0183 (0.0092); δ: 0.0151 (0.0085)].

Finally, to investigate the robustness of our results to the use of coverage ratios in one more way, we reestimated the model allowing for two different protection equations, one where the dependent variable included only price-oriented nontariff barriers (NTBs) (antidumping duties, countervailing duties) and one where it included quantity measures (quotas, VERs, etc.). The results from both specifications were very similar. Moreover, the magnitude of the parameters was quite intuitive; using only price- (or quantity-) oriented NTBs as a measure of protection produced estimates of the structural parameters β and α_L that were higher than the ones obtained using the composite measure. Intuitively, including only one component of nontariff protection in the estimation ignores a significant fraction of protection, overestimating the "openness" of the economy; the implied welfare weight is then higher, and so is the estimated degree of lobby representation.

IV. Conclusion

In this paper we depart from the existing empirical literature on endogenous trade policy, in that we let our estimation be closely

[21] Equation (4) becomes:

$$\frac{t_i^*}{1 + t_i^*} e_i \frac{M_i}{X_i} = \gamma + \delta I_i + \epsilon_i$$

where γ now represents a constant. The estimating model consists of the above equation and equation (7); the import-penetration equation (6) is no more estimated.

guided by a theoretical model. We see three major benefits of this approach. First, of course, this allows a rigorous empirical examination of the model at hand. Second, it allows one to estimate structural parameters, which can contain a great deal of information. Third, we think there is value in utilizing a parsimonious model. The spirit of our research has been to look for the "minimal efficient" model among a number of alternative specifications, in the sense of a model that predicts trade protection in the most accurate way with the simplest, theoretically sound specification. The strict version of the G-H model seems not too far from satisfying these requirements. In particular, the qualitative predictions of the G-H model are consistent with our data, and none of the additional variables commonly employed in previous studies (such as unionization, concentration, changes in import penetration, etc.) substantially improves the explanatory power of the strict G-H model.

While one cannot claim full empirical success for the model, since some of its predictions find only weak support, the mere fact that it is not inconsistent with our data is remarkable. Tests of the strict versions of trade models traditionally yield disastrous results for the theories under investigation—the poor empirical performance of the strict version of the Heckscher-Ohlin model is a good example. It takes several extensions and modifications before such models can begin to fit the data. For this reason, we were surprised to find that the strict version of the G-H model is not "grossly" inconsistent with the data.

Adopting a theory-guided approach is important not only because the theoretical model suggests what variables should be included in the regressions and which are endogenous or exogenous, but also because it suggests the way these variables interact with each other, that is, the functional forms that should be utilized. This seems to affect the conclusions about the determination of trade policies. In particular, using the interactive specification suggested by the model, we find that the positive correlation between protection and import penetration documented in previous empirical studies applies only to the group of nonorganized sectors. Within the group of politically organized sectors, the point estimate suggests—in line with the theoretical predictions—a negative correlation between the two variables. Even though the latter estimate is not statistically significant, there seems to be a clear difference in the pattern of protection received by the two groups. It is interesting to note that this difference goes in the direction predicted by the model: it is the nonorganized sectors that exhibit the positive correlation between import penetration and protection. For politically organized industries, higher import penetration implies lower protection.

Of some interest, we believe, is also our estimate of a key structural parameter of the model, namely the weight attached by the government to social welfare. We estimate this weight to be many times higher than the weight attached to contributions. Overall, then, our results suggest that the G-H model is consistent with our data and helps explain the cross-sectoral structure of trade protection, but that the magnitude of political considerations in the government's objective is small.

Appendix A: Results from Reduced-Form Equations

Table A1—Reduced-Form Equation for Import Penetration

Variable	Coefficient	Standard error	t
Physical capital	1296.03	435.28	2.98
Inventories	−2501.81	709.42	−3.53
Engineers and scientists	−52.79	405.38	−0.13
White-collar	489.58	192.54	2.54
Skilled	91.86	252.80	0.36
Semiskilled	−99.89	202.18	−0.49
Cropland	514.86	231.45	2.23
Pasture	−1493.13	366.34	−4.08
Forest	4262.02	3451.79	1.24
Coal	3043.31	3356.99	0.91
Petroleum	5.62	154.04	0.04
Minerals	−4778.79	4020.54	−1.19
Seller concentration	−37.46	56.29	−0.67
Seller number of firms	−5.64	28.03	−0.20
Buyer concentration	−78.17	114.22	−0.68
Buyer number of firms	−2.35	12.76	−0.18
Scale	−10.80	249.05	−0.04
Capital stock	−118.88	42.99	−2.77
Unionization	79.64	46.55	1.71
Geographic concentration	−43.02	44.57	−0.97
Tenure	−0.21	4.73	−0.05
Constant	−11.62	79.70	−0.15

Notes: Dependent variable: Inverse penetration ratio (X_i/M_i).
Number of observations: 107.
R^2: 0.49.

Table A2—Reduced-Form Equation for Political Organization

Variable	Coefficient	Standard error	t
Physical capital	23.90	15.25	1.57
Inventories	−5.14	33.83	−0.15
Engineers and scientists	21.03	15.73	1.34
White-collar	−7.30	6.71	−1.09
Skilled	22.34	10.97	2.04
Semiskilled	−18.53	7.67	−2.42
Cropland	15.57	21.13	0.74
Pasture	−25.88	35.86	−0.72
Forest	−11.87	123.90	0.10
Coal	136.72	104.36	1.31
Petroleum	10.45	6.06	1.73
Minerals	63.35	214.73	0.30
Seller concentration	−3.32	1.76	−1.88
Seller number of firms	−1.69	1.23	−1.38
Buyer concentration	2.13	4.27	0.50
Buyer number of firms	−0.53	0.56	−0.96
Scale	50.81	18.84	2.70
Capital stock	−2.83	1.51	−1.88
Unionization	2.16	1.58	1.37
Geographic concentration	−3.22	1.62	−1.99
Tenure	−0.44	0.16	−2.73
Constant	2.59	2.92	0.89

Notes: Dependent variable: Political-organization dummy (I_i).
Number of observations: 107.

Table A3—Results from an Extended Specification ($\mu = 1$)

Variable	Coefficient	Standard error	t
X_i/M_i	−0.0092	0.0044	−2.104
$(X_i/M_i) * I_i$	0.0089	0.0089	1.998
Constant	−0.2545	0.2409	−1.057
I_i	0.3851	0.3466	1.111

Note: Dependent variable: ($t_i^* e_i/1 + t_i^*$).

APPENDIX B: DEFINITION OF ORGANIZATION DUMMY

TABLE B1—LIST OF UNORGANIZED SECTORS

Code	Standard Industrial Classification (SIC)	Code	Standard Industrial Classification (SIC)
SIC 315	Leather Gloves and Mittens	SIC 227	Floor Covering Mills
SIC 319	Leather Goods, Not Elsewhere Classified	SIC 228	Yarn and Thread Mills
SIC 316	Luggage	SIC 321	Flat Glass
SIC 313	Boot and Shoe Cut Stock and Findings	SIC 328	Cut Stone and Stone Products
SIC 311	Leather Tanning and Finishing	SIC 241	Logging Camps and Logging Contractors
SIC 213	Tobacco (Chewing and Smoking) and Snuff	SIC 251	Household Furniture
SIC 212	Cigars	SIC 324	Cement, Hydraulic
SIC 314	Footwear, Except Rubber	SIC 322	Glass and Glassware, Pressed or Blown
SIC 317	Handbags and Other Personal Leather Goods	SIC 306	Fabricated Rubber Products, Not Elsewhere Classified
SIC 393	Musical Instruments	SIC 301	Tires and Inner Tubes
SIC 231	Men's, Youths', and Boys' Suits, Coats, and Overcoats	SIC 302	Rubber and Plastics Footwear
		SIC 385	Ophthalmic Goods
SIC 259	Miscellaneous Furniture and Fixtures	SIC 383	Optical Instruments and Lenses
SIC 278	Blankbooks, Looseleaf Binders, and Bookbinding and Related Work	SIC 387	Watches, Clocks, Clockwork Operated Devices, and Parts
SIC 273	Books	SIC 381	Engineering, Laboratory, Scientific, and Research Instruments, and Associated Equipment
SIC 271	Newspapers: Publishing, Publishing and Printing		
SIC 272	Periodicals: Publishing, Publishing and Printing	SIC 386	Photographic Equipment and Supplies
SIC 399	Miscellaneous Manufacturing Industries	SIC 339	Miscellaneous Primary Metal Products
SIC 395	Pens, Pencils, and Other Office and Artists' Materials	SIC 332	Iron and Steel Founderies
SIC 394	Toys and Amusement, Sporting, and Athletic Goods	SIC 249	Miscellaneous Wood Products
SIC 391	Jewelry, Silverware, and Plated Ware	SIC 244	Wood Containers
SIC 275	Commercial Printing	SIC 307	Miscellaneous Plastics Products
SIC 396	Costume Jewelry, Costume Novelties, Buttons, and Miscellaneous Notions, Except Precious Metal	SIC 295	Paving and Roofing Materials

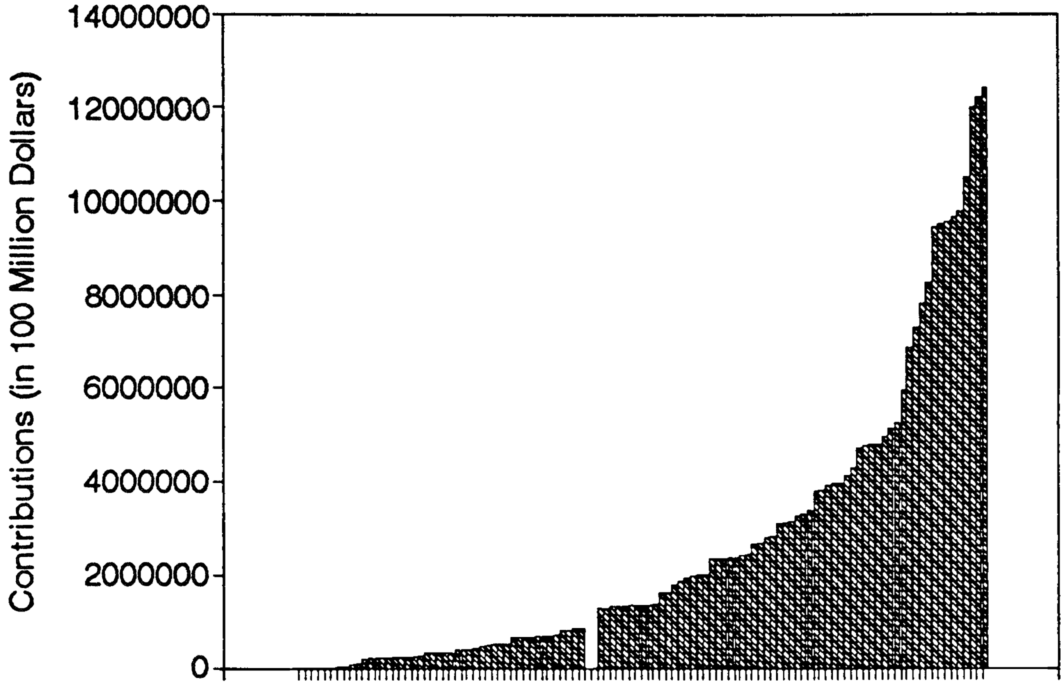

FIGURE B1. BAR CHART OF PAC CONTRIBUTIONS BY SIC3

Appendix C: Testing for Heteroskedasticity

The test for heteroskedasticity we report in Section III, subsection B, is based on a simplified version of the econometric model. This version has $t_i^* e_i$ on the left-hand side, and the estimation results are very similar to the ones we obtain by estimating the specification (4)–(8). Because of the complexity of the formulas involved in the computation of the test statistic, we also decided to ignore the censoring at $t_i = 1$, given that there is only one observation at 1 in our data.

Let us write the modified version of equation (4) as follows:

$$y_i^* = X_i'\beta + \epsilon_i \tag{4'}$$

where $y_i^* = t_i^* e_i$ and $X_i'\beta = \gamma(X_i/M_i) + \delta I_i(X_i/M_i)$. The Tobit corresponding to this specification is:

$$y_i = \begin{cases} y_i^* & \text{if } y_i^* > 0 \\ 0 & \text{if } y_i^* \leq 0 \end{cases} . \tag{5'}$$

The rest of the model [that is, equations (6)–(8)] is the same as before.

To test for heteroskedasticity, we conduct a test similar to the one described in Chesher and Irish (1987 pp. 35–43) and Adrian Pagan and Frank Vella (1989 pp. 35–38). The basic idea of the test is to test for correlations between moments of the model residuals and variables that may be related to the variance of the error term. In the presence of censoring, conventional residuals cannot be calculated; however, it is still possible to use an alternative definition of residuals (what we call "generalized residuals"), to produce estimates of these correlations.

Our first step is to calculate the generalized residuals corresponding to model (4′)–(8). Note that there are three cases to consider. In the first case, the dependent variable is observed. Then the residuals can be calculated in the usual way. The second case corresponds to a censored dependent variable in (4′), and a positive observation in the probit model of equation (7). Finally, the third case is the one where the dependent variable in (4′) is censored, and the probit variable in (7) is zero. Hence, the generalized residuals take the form:

$$\hat{\epsilon}_i^2 = \begin{cases} (y_i - X_i'\beta)^2 \\ \quad \text{if } y_i^* \text{ is observed, that is } y_i = y_i^* \\ E(\epsilon_i^2 | \epsilon_i < -X_i'\beta, u_{2i} > -\zeta_2' \mathbf{Z}_{2i}, u_{1i} = c) \\ \quad \text{if } y_i^* \leq 0 \text{ and } I_i^* > 0 \\ E(\epsilon_i^2 | \epsilon_i < -X_i'\beta, u_{2i} < -\zeta_2' \mathbf{Z}_{2i}, u_{1i} = c) \\ \quad \text{if } y_i^* \leq 0 \text{ and } I_i^* \leq 0. \end{cases}$$

To compute the expectation terms corresponding to the last two cases, we first compute the conditional bivariate normal distribution $(\epsilon, u_2 | u_1)$, and then utilize the formulas for the moments of the truncated bivariate normal distribution that are provided in G. S. Maddala (1986 p. 368). As evident from these formulas, the computation of the expectation terms is quite cumbersome, even in the simplified version of the model.

Once the generalized residuals are calculated, we form the vector $V_i = (\hat{\epsilon}_i^2 - \hat{\sigma}^2)\mathbf{Z}_i$, where $\hat{\sigma}^2$ is our estimate of the variance of the error term, and $\mathbf{Z}_i$ consists of the variables we suspect are correlated with the variance of ϵ. In our case, there are two obvious candidates for $\mathbf{Z}_i$: the elasticities, and their standard errors as reported in Shiells et al. (1986).

The final step in the test involves running the pseudo-regression

$$\mathbb{i} = RC + u$$

where $\mathbb{i}$ is a $(N \times 1)$ vector of ones (N is the number of observations), R includes V and the $(\mathbf{N} \times \mathbf{K})$ matrix of the gradients of the likelihood function with respect to the parameters of the model. The explained sum of squared residuals from OLS estimation of the above regression is our test statistic; it is distributed as χ^2, with degrees of freedom equal to the number of restrictions.

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