"PROTECTION FOR SALE" IN A DEVELOPING COUNTRY: DEMOCRACY VS. DICTATORSHIP

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Abstract—For a "genuine" small open economy that has experienced both dictatorship and democracy, we find support for the predictions of the Grossman-Helpman (1994) "Protection for Sale" model. In contrast to previous studies, we use various protection measures (including tariffs, the direct measure of the theoretical model) and perform both single-year and panel regressions. Using Turkish industry-level data, the government's weight on welfare is estimated to be much larger than that on contributions. More importantly, we find that this weight is generally higher for the democratic regime than for dictatorship.

I. Introduction

THE literature on the political economy of trade policy has evolved over the last two decades into two strands, one focusing on majority voting and the other emphasizing special-interest politics. The majority voting approach to trade policy was introduced by Mayer (1984),[1] and the pioneering model in the special-interest strand is Findlay and Wellisz (1982), where the tariff itself is assumed to be an exogenously given function of resources into lobbying by different types of factor owners.[2] Electoral competition was introduced into lobbying models by Magee, Brock, and Young (1989), and Hillman (1989) provided an altogether different approach to modeling special-interest politics by introducing the idea of the "political support" function (which the government maximizes) that incorporates the government's preferential treatment of an organized industry as well as the cost of protecting this industry given by the excess burden on society.

Following this path, the special interest literature on trade policy has evolved into the state-of-the-art Grossman and Helpman (1994) "Protection for Sale" model.[3] This model is path-breaking for several reasons. Firstly, its framework is multisectoral. Secondly, it provides microfoundations to the behavior of lobbies and politicians. The government's objective function is a weighted sum of political contributions and aggregate welfare, and each lobby maximizes its welfare net of political contributions. Most importantly, the level of protection is derived as an estimable function of industry characteristics and other political and economic factors.

Two recent empirical papers, Goldberg and Maggi (1999) and Gawande and Bandyopadhyay (2000) estimate the Grossman-Helpman protection expressions using industry-level data from the United States.[4,5] The two papers are similar in the questions they address, but they are somewhat different in the details of their approaches. Whereas Goldberg and Maggi restrict focus on the protection expressions, Gawande and Bandyopadhyay concentrate more on the lobbying aspects and the determinants of the magnitude of contributions. Goldberg and Maggi use the basic Grossman-Helpman framework, whereas Gawande and Bandopadhyay introduce intermediate goods. The econometric specifications are somewhat different in the two papers. However, the results from both these papers are very similar in that they find that the weight on aggregate welfare in the government's objective function is several times higher than that on contributions. As predicted by the Grossman-Helpman model, both papers find that protection to organized sectors is negatively related to import penetration and the (absolute value of) import demand elasticity, while protection to unorganized sectors is positively related to these two variables.

We investigate these Grossman-Helpman predictions using industry-level data from Turkey. Our paper differs from the existing papers in the literature in the following respects. (i) We look at the cross-industry protection levels in a developing country for four different years in the period 1983 to 1990 (as opposed to a developed country for a single year). (ii) Our data set spans both dictatorial and democratic regimes. (iii) We use a variety of protection measures: nominal protection rates (the direct measure suggested by the theoretical model), effective protection rates

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[1] The Mayer (1984) framework has been extended to the issue of the choice of instruments in a series of papers by Mayer and Riezman (1987, 1989, 1990).

[2] Papers using a similar approach are Feenstra and Bhagwati (1982) and Rodrik (1986).

[3] The basic framework of Grossman and Helpman (1994) has been used to analyze issues of trade negotiations and free trade areas in Grossman and Helpman (1995a, 1995b), commitment to trade agreements by Maggi and Rodriguez-Clare (1998), issues of lobby formation by Mitra (1999), and the relationship between lobby formation and reciprocity in trade policy in Krishna and Mitra (2000).

[4] There is a recent unpublished paper by McCalman (2000) that empirically investigates the "Protection for Sale" model for Australia for two separate years (1968–1969 and 1991–1992) and, unlike the two papers on the United States, focuses on tariffs (and not on NTB coverage ratios). Again, McCalman finds the model to be consistent with the data. He is also able to analyze the endogeneity of the Australian trade liberalization and attributes it to the increase in the fraction of the population that is politically organized and the government's weight on welfare relative to contributions.

[5] Other empirical papers investigating endogenous trade policy issues (but outside the Grossman-Helpman framework) are Baldwin (1985), Trefler (1993), Ray (1981), and Gawande (1997a, 1997b, 1998).

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and NTB coverage ratios (the measure used in the previous studies on the United States). (iv) We perform cross-sectional analysis for each year separately as well as panel regressions for the entire time span.

We find strong support for the validity of the fundamental predictions of the Grossman-Helpman "Protection for Sale" model. As in the previous studies, the government's weight on welfare is estimated to be much larger than that on contributions. Additionally, this weight on welfare is higher for the democratic regime than for dictatorship.

In this context, it is useful to discuss the rationale for our particular choice of country. Firstly, the Grossman-Helpman (1994) "Protection for Sale" model makes a "small, open economy" assumption that is certainly more apt for Turkey than for the United States.

Secondly, the use of Turkish data can tell us something about the applicability of this model to a developing country. In almost any country, interest-group contributions (although well documented only in the United States) are an integral part of the political environment. They are made mainly to finance election campaigns or take the form of bribes to politicians and other key government officials. In Turkey, there have been newspaper reports of well-known industrialists contributing in various forms to election campaigns.[6] Such support has included even the provision of personal airplanes and helicopters.[7] Further, in local newspaper reports, we find allegations of politicians accepting bribes in exchange for help in obtaining government contracts.

Going back to the choice of country, the third issue is the applicability of the "Protection for Sale" framework across different political systems. Turkey has had both dictatorial and democratic regimes and our data set spans both these regimes. In a democracy, on the one hand, there is need for political contributions as governments need to spend on advertising (campaigns) to get reelected. On the other hand, there is a cost of receiving political contributions because in exchange the government has to provide preferential treatment to the donors even if that leads to a reduction in average welfare or the quality of life. In dictatorships, the need for such contributions is lower, but the affinity for contributions (on the part of the ruler) may still exist.[8] Additionally, the dictator may not need to care as much about the contributions–welfare tradeoff. Therefore, what kind of political system would assign a greater weight to welfare relative to contributions is an important empirical question.

Extensive data on different measures of protection, import penetration, import-demand elasticities, and so on are available for Turkey. Because no data were available on trade-related (or other) political contributions, we determine whether sectors are politically organized from the membership data for the Turkish Industrialists and Businessmen Association (TUSIAD) and then statistically validate this determination using classification methods based on discriminant analysis.

Besides the use of classification methods in constructing the political organization variable, our paper features a few other methodological advances. We depart from the estimated linear specifications of the previous studies and directly estimate, using nonlinear methods, the parameters (and their standard errors) in the structural equation of interest. The availability of data on a large number of exogenous (instrumental) variables enables us to perform nonlinear 2SLS. Allowing for year-specific effects, we also estimate (by both nonlinear 2SLS as well as generalized method of moments (GMM)) our structural parameters for the data pooled across years. Some hypotheses of economic interest, involving the structural parameters, are also tested.

In section II, we discuss the theoretical framework. The econometric methodology is presented in section III. Section IV describes the data and their sources. In section V, we present the results of our estimation. Finally, section VI concludes.

II. Theoretical Framework

In this section, we provide an abridged description of the "Protection for Sale" model of Grossman and Helpman (1994).

Consider a small, open economy. Individuals are assumed to have identical preferences. Each individual possesses labor and at most one kind of specific factor of production. There are N nonnumeraire goods, each requiring a different kind of factor of production specific to that good and labor. In addition, there is a numeraire good that is produced under constant returns to scale using only labor. A quasi-linear utility function (linear in numeraire good consumption, concave in the consumption of each nonnumeraire good, and additively separable across all goods) is assumed.

It is also assumed that the only policy instruments available to politicians are trade taxes and subsidies. Further, the government redistributes revenue uniformly to all its citizens.

In sectors that are politically organized, the specific-factor owners are able to lobby the government for prefer-

[6] For example, the daily newspaper *Milliyet* had a report on September 3, 1991, about some top businessmen (Kamhi, Boyner, Ekinci, Alaton, and Bodur) providing substantial support in different forms for the election campaign of the Motherland-True Path Party coalition.

[7] For instance, there are reports that Cavit Caglar, a well-known textile producer, provided an airplane and two helicopters for the election campaign of Suleyman Demirel, the president of the True Path Party.

[8] Grossman and Helpman (1994) write "Organized interest groups are able to offer political contributions, which politicians value for their potential use in the coming election (and perhaps otherwise)" (p. 834). It is this "perhaps otherwise" component that might be fairly important in dictatorships, and dictators might, for their own benefit, want to auction off policies. Often dictators also might rely on the support of politically and economically powerful individuals and groups for their existence. In return, they have to provide these powerful individuals or groups with different kinds of concessions. Thus, Grossman and Helpman (1994) write "Such an objective function seems plausible for a government that is concerned about the next election, but *broader interpretations* [emphasis added] also are possible" (p. 836).

ential treatment in the form of higher trade protection for their own sectors and lower protection for other sectors. The interaction between the government and the lobbies takes the form of a "menu auction" as in Bernheim and Whinston (1986). Thus, we have the following two-stage game.

In the first stage, lobbies provide the government with their contribution schedules taking into account the government's objective function (described later). Each lobby takes the contribution schedules of other lobbies as given.

In the second stage, taking into account the contribution or offer schedules from the previous stage, the government sets trade policy to maximize a weighted sum of political contributions and overall social welfare.

The government's objective function is

$$\Omega_G(p) = \sum_{j \in \Lambda} C_j(p) + a\Omega_A(p), \tag{1}$$

where Λ is the set of organized interest groups (lobbies),
$p \in P$ is the domestic price vector,
$\Omega_A(p)$ is aggregate social welfare,
$C_j(p)$ is the contribution schedule of the jth lobby and,
P is the set of domestic price vectors from which the government may choose.

The set P is bounded such that each domestic price lies between some minimum and some maximum value. Grossman and Helpman (1994) restrict attention to equilibria that lie in the interior of P. The parameter a in equation (1) is the weight that the government attaches to aggregate social welfare relative to political contributions. The higher a, the lower is the government's affinity for political contributions and the higher is its concern for social welfare.

Grossman and Helpman show that, with contribution schedules that are continuous in the price vector in the neighborhood of the equilibrium, the government's problem of choosing its most preferred tariff vector (on receiving all the contribution schedules) is equivalent to maximizing the following function with respect to the domestic price vector p:

$$\sum_{j \in \Lambda} \Omega_j(p) + a\Omega_A(p), \tag{2}$$

where $\Omega_j(p)$ is welfare of sector j.[9] This maximization yields trade taxes and subsidies that satisfy the following, familiar Grossman-Helpman modified "Ramsey Rule":

$$\frac{t_i}{1+t_i} = \frac{I_i - \alpha_L}{a + \alpha_L} \cdot \frac{z_i}{e_i} \tag{3}$$

where z_i is the ratio of domestic output to imports or exports (depending on whether the sector is import competing or an exporting one),
e_i is the absolute value of price elasticity of import demand or export supply,
I_i is an indicator variable that takes a value 1 if the sector is politically organized and 0 for an unorganized sector, and
α_L is the proportion of the country's population that is organized.

This tariff expression can be written as the following empirically estimable form:

$$\frac{t_i}{1+t_i} = \frac{1}{a + \alpha_L}\left[I_i \cdot \left(\frac{z_i}{e_i}\right)\right] - \frac{\alpha_L}{a + \alpha_L}\left[\frac{z_i}{e_i}\right] \tag{4}$$

Because $a \in [0, \infty)$ and $\alpha_L \in [0, 1]$, the coefficient of $[I_i \cdot (z_i/e_i)]$ should be positive and that of (z_i/e_i) should be negative. Also, the coefficient of the former is larger in magnitude than the latter. This means that the protection is increasing in (z_i/e_i) for an organized industry, but decreasing in (z_i/e_i) in the case of an unorganized sector. In fact, in the theoretical model, organized sectors are given positive protection, whereas unorganized sectors are exploited through negative protection (which will be modified in the empirical model through the use of a constant term). Thus, there is deviation from free trade in opposite directions for organized and unorganized sectors, and the size of this deviation is increasing in (z_i/e_i) because (i) the deadweight costs of this deviation are increasing in the magnitude of the trade elasticities, (ii) the benefits to lobbies from protection are higher if their output levels are higher, and (iii) the costs of deviation from free trade are lower, the lower is the volume of actual and potential trade.

Because, in Turkish agriculture, median voter as well as political support concerns other than contributions could be important in determining tariffs, we also experiment with the following specification:

$$\frac{t_i}{1+t_i} = \frac{1}{a + \alpha_L}\left[I_i \cdot \left(\frac{z_i}{e_i}\right)\right] - \frac{\alpha_L}{a + \alpha_L}\left[\frac{z_i}{e_i}\right] + \beta\left[I_{Ai} \cdot \left(\frac{z_i}{e_i}\right)\right], \tag{5}$$

where I_{Ai} takes the value 1 if sector i is agricultural and 0 otherwise. β, thus, captures the populist protection to agriculture, which was viewed as a large source of votes.[10] Further, this specification takes care of mistakes made in our judgment of whether agricultural sectors are organized.

[9] This new reduced-form maximand is an additively separable form of the more general Hillman (1989) political support function; that is, Grossman and Helpman (1994) provide microfoundations for models that use the political support function approach. Alternatively, in Goldberg and Maggi (1999) and Maggi and Rodriguez-Clare (1998), the interaction between the government and lobbies is modeled as a Nash bargaining game over trade policy and contributions.

[10] Helpman (1995) shows how median-voter and political-support forces can affect tariffs within the kind of multisectoral, specific-factors framework presented in this paper. He shows that $\left(\frac{z_i}{e_i}\right)$ is still the main determinant of a sector's tariff under those forces.

III. Econometric Methodology

We deviate from Goldberg and Maggi (1999) and Gawande and Bandyopadhyay (2000) in that we directly estimate the structural coefficients and their standard errors by putting the estimation problem in its appropriate nonlinear regression form.[11] Although the linear regressions considered in the previous papers are appealing and straightforward to implement, the additional computational cost of directly estimating the structural coefficients is minimal. Moreover, we can immediately obtain the appropriate standard errors for the structural coefficients.

In proceeding, we are faced with two alternatives: we could either keep the elasticity in the tariff expression on the right-hand side, as in Gawande and Bandyopadhyay, or transfer it to the left-hand side, as in Goldberg and Maggi. For the Turkish data, we feel we are better off with the import-demand elasticity on the right-hand side for the following two reasons:

- Our elasticity estimates are estimated with much greater precision (lower measurement error), almost all of them being significant at the 1% through 5% significance levels. In addition, the presence of any remaining measurement error is handled through the use of instrumental variables in estimation, as explained later.
- Only when we keep the elasticity on the right-hand side do we have a consistent set of instrumental variables as we move from one year to another.

We next give a brief description of the model we estimate in this paper.

A. Single Year Estimation

Let $\theta_j = (c_j, \alpha_{Lj}, a_j)'$ be the (3×1) vector of structural parameters of the model for year j, where c_j is the constant term. Let $y_{ij} = t_{ij}/(1 + t_{ij})$ denote the left-side variable, for sector i and year j, and $x_{ij} = \{1, [I_i \cdot (z_{ij}/e_{ij})], (z_{ij}/e_{ij})\}$ be the (1×3) vector of explanatory variables. We denote by $g(x_{ij}; \theta)$ the right-side nonlinear function

$$g(x_{ij}, \theta_j) = c_j + \frac{1}{a_j + \alpha_{Lj}}\left[I_i \cdot \left(\frac{z_{ij}}{e_{ij}}\right)\right] - \frac{\alpha_{Lj}}{a_j + \alpha_{Lj}}\left[\frac{z_{ij}}{e_{ij}}\right] \tag{6}$$

so that the estimable equation has a standard nonlinear regression format with additive errors:

$$y_{ij} = g(x_{ij}, \theta_j) + u_{ij}. \tag{7}$$

$\{u_{ij}\}_{i=1}^{n=37}$ is the regression error term that captures possible measurement errors in the right-side variables, as well as other factors (outside the theoretical model) that affect the determination of the tariff.

Equation (7) cannot be estimated by nonlinear least squares because the right-side variables may be correlated with the regression error term because of the endogeneity of our right-side variables (z, e, and I) with respect to the tariff (see Goldberg and Maggi (1999), Mitra (1999), and Gawande and Bandyopadhyay (2000)), and also because of the measurement error associated with the political organization variable, I (from the possibility of misclassification) and the import demand elasticity, e (because it is estimated). Both the problem of endogeneity and measurement error of the right-side variables can be handled by using instrumental variables. These should be variables that are correlated with the right-side variables themselves but not correlated with the regression error. Shifts and rotations of the import demand function will affect both the import demand elasticity and the import penetration ratio, at any given tariff. Thus, we try to identify such shift variables affecting import demand. These variables include different forms of domestic (nontrade) governmental concessions, sector-specific minimum wages, incentives to freight, and so on. As regards the organization indicator, I, the degree of unionization may be an important determinant of the political organization of skilled workers (owners of sector-specific human capital) in any sector. Besides, as lobby formation costs are sunk in nature, import growth (by helping the formation of expectations regarding the future intensity of import competition) may also determine I.[12] It is important to note in this context that no existing theory suggests that any of the preceding instrumental variables are either endogenous to the tariff or are the determinants of the tariff missing from the right-side of the regression equation (7). Also, these instruments are correlated with the true values of the right-side variables (and are not expected to be correlated with any of the measurement errors). Finally, and very importantly, after estimation we perform the appropriate test for the validity of our instruments (see following).

Thus, the preceding regression is estimated by nonlinear two-stage least squares (NL2SLS)[13] so as to retrieve an estimator of θ_j directly, say $\hat{\theta}_{jn}$. Let y_j denote the $(n\times 1)$ vector of observations of the left-side variable, X_j denote the $(n\times 3)$ matrix of observations of the right-side variables, and W_j denote the $(n\times k)$ matrix of instrumental variables.[14]

[11] As can be seen from equation (4), the right-side expression is nonlinear in both variables and parameters.

[12] Decisions to get politically organized are based on the expected magnitude of protection on being organized relative to that on remaining unorganized. Thus, because z and e determine protection, the determinants of z and e will also determine I.

[13] See chapter 7 and 17 of Davidson and McKinnon (1993) or chapter 6 of Dhrymes (1994) for textbook discussions on estimation and testing of nonlinear regression models using NL2SLS and GMM.

[14] Standard practice in NL2SLS does not restrict the list of instrumental variables used in estimation to the levels of the instruments, but permits the use of their squares and cross-products as well. We follow this practice: using a small set of exogenous variables as instruments, almost exclusively levels and squares, is sufficient for producing economically meaningful results across all years.

Finally, let $\Theta \subseteq R^3$ denote the admissible parameter space. The estimator $\hat{\theta}_{jn}$ is then obtained as

$$Q_n(y_j, X_j; \hat{\theta}_{jn}) = \inf_{\theta \in \Theta} \{[y_j - g(X_j; \theta_j)]' W_j (W_j' W_j)^{-1} \times W_j' [y_j - g(X_j; \theta_j)]\}. \tag{8}$$

The corresponding standard errors of $\hat{\theta}_{jn}$ are obtained from the diagonal elements of the estimated covariance matrix of $\hat{\theta}_{jn}$,

$$\text{Est. Cov } (\hat{\theta}_{jn}) = \hat{\sigma}_j^2 G_j(\hat{\theta}_{jn})' W_j (W_j' W_j)^{-1} W_j' G_j(\hat{\theta}_{jn}), \tag{9}$$

where $\hat{\sigma}_j^2$ is the estimated residual variance and $G_j(\hat{\theta}_{jn})$ is the $(n \times 3)$ Jacobian matrix of $g(X_j; \theta_j)$ evaluated at $\hat{\theta}_{jn}$.

With enough instruments at our disposal ($k >$ # parameters to be estimated = 3), all equations are overidentified. Thus, after estimation, we test whether the $k - 3$ additional instrumental variables are correlated with the regression error term. The (overidentification) test statistic is given by nR_u^2, where R_u^2 is the coefficient of determination of an auxilliary LS regression of the estimated NL2SLS residuals $\hat{u}_{ij} = y_{ij} - g(x_{ij}, \hat{\theta}_{jn})$ on the matrix of instrumental variables W_j. The statistic is distributed as χ^2_{k-3}.[15]

Some hypotheses of interest were tested after estimation. We summarize these hypotheses here.

1. $H_0(1) : 1/a_j = 0$, an approximate test of the hypothesis that the government maximizes aggregate welfare and that the composite coefficient of $I_i(z_{ij}/e_{ij})$ is significant. Note that this hypothesis is equivalent to the hypothesis $H_0'(1) : 1/(a_j + \alpha_{Lj}) = 0$, contingent on α_{Lj} lying in the relevant economic range of [0, 1].[16]
2. $H_0(2) : \alpha_L/(a_j + \alpha_{Lj}) = 0$, a test for the hypothesis that the composite coefficient of (z_{ij}/e_{ij}) is significant. Note that this test is affected by both the value of α_{Lj} and a_j; a large a_j will tend to reduce the value of the composite coefficient independently of the value of α_{Lj}.
3. $H_0(3) : \alpha_{Lj} = 0$ and $1/a_j = 0$, a (joint) test for the significance of the model. This is the most appropriate test in examining the joint significance of the composite coefficients of (z_{ij}/e_{ij}) and $I_i(z_{ij}/e_{ij})$.

[15] The acceptance of the null hypothesis effectively supports lack of correlation of any of the instrumental variables with the regression error.

[16] Note that the distribution of the relevant test statistic is well defined. What is being tested is whether a_j is large enough so that its reciprocal $1/a_j$ is statistically indistinguishable from 0 (even though this reciprocal cannot take the value 0 for any finite a_j).

All hypotheses involve nonlinear restrictions and were tested using a Wald-type test statistic.[17]

B. *Panel Estimation*

As a final step in our empirical analysis, we estimated the specification of equation (7) by pooling our data across years and including year dummy variables using a common list of instruments for all years, presented as NL2SLS(DV) in our tables, and by using panel fixed-effects NL2SLS and GMM estimation with year-specific instruments, to account both for year-specific intercepts and year-specific error variances (presented as NL2SLS (FE-SI) and GMM (FE-SI), respectively).[18]

We define the vector of year dummies, D_{ij}, with corresponding coefficient vector, δ. We excluded 1983 (the dictatorship year) from the dummy variable list so that the dummy coefficients represent contrasts with respect to that year. Let $\beta = (\theta', \delta')'$ denote the new coefficient vector. Equation (7) can now be rewritten as

$$y_{ij} = g^*(x_{ij}, D_{ij}; \beta) + u_{ij}, \tag{10}$$

where $g^*(x_{ij}, D_{ij}; \beta) = g(x_{ij}; \theta) + D_{ij}\delta$.

Estimation of equation (10) by pooled NL2SLS (DV) is based on the same objective function as in equation (8) by defining W to denote the common instrumental variables matrix. For the panel NL2SLS (FE-SI), estimation is based on the same objective function as in equation (8), but we now define W as the block diagonal matrix $W = \text{diag}(W_{1983}, W_{1984}, W_{1988}, W_{1990})$ and we substitute (y, X) for (y_j, X_j), where $y = (y'_{1983}, y'_{1984}, y'_{1988}, y'_{1990})'$ and $X = (X'_{1983}, X'_{1984}, X'_{1988}, X'_{1990})'$. Letting $B \subseteq R^6$ denote the new admissible parameter space, the panel-GMM estimator, $\hat{\beta}_n$ is obtained as

$$Q_n^*(y, X, D; \hat{\beta}_n) = \inf_{\beta \in B} \{[y - g^*(X, D; \beta)]' W V^{-1} \times W' [y - g^*(X, D; \beta)]\}, \tag{11}$$

where now W is as in the case of panel NL2SLS (FE-SI).[19]

[17] Let $R(\theta_j) = 0$ denote the nonlinear function imposing J restrictions, and let $r(\theta_j) = \partial R(\theta_j)/\partial \theta_j$. The actual test statistic takes the following form:

$$\mathcal{W}_{jn} = (n/J) R(\hat{\theta}_{jn})' [r(\hat{\theta}_{jn}) \text{Est. Cov } (\hat{\theta}_{jn}) r(\hat{\theta}_{jn})']^{-1} R(\hat{\theta}_{jn}),$$

whose values we can compare to the critical values from the $F_{(J,n-3)}$ distribution. For a single restriction $J = 1$, we have that $\sqrt{\mathcal{W}_{jn}}$ has an asymptotic standard normal distribution.

[18] The list of instruments used in estimating the pooled equation with dummy variables includes the (union of the) exogenous variables and their squares used in single-year equations as well as some cross products of these variables, pooled across years. On the other hand, the list of instruments used in estimating the other panel equations include the year-specific instrumental variables used in single-year equations.

[19] In the case of FE-SI regressions, we report the coefficients of four year-specific constants instead of a constant and the coefficients on three year dummy variables, as in the case of the DV regressions.

The weighting matrix V is the covariance of the moment restriction; that is,

$$V = E(W'[y - g^*(X, D; \beta)][y - g^*(X, D; \beta)]'W),$$

where we use White's heteroskedastic covariance matrix estimator in obtaining an estimate of V, to account for different error variances, σ_j^2, across years.

IV. Data

We need data on imports and output to calculate the import penetration ratio. Data on imports are obtained from the *UN International Trade Statistics Yearbook* (various issues), and those on domestic output are obtained from the *U.N. Statistical Yearbook* (various issues), *Monthly Bulletin of Statistics* (various issues), the web site of the *U.N. Food and Agriculture Organization* (http://www.fao.org), *OECD Industrial Structure Statistics* (various issues), and *U.N. Industrial Statistics Yearbook* (various issues).[20]

Import-demand elasticities for thirteen of the 37 product categories in our study are directly obtained from Thomakos and Ulubaşoğlu (2000), who followed the methodology of Shiells, Stern, and Deardorff (1986). The remaining 24 elasticities were estimated for this study with the same techniques as in Thomakos and Ulubaşoğlu.

Data on nominal rates of protection (NRP), effective rates of protection (ERP), and non-tariff barrier coverage ratios (NTB) are obtained from Togan (1994).[21]

As explained in subsection IIIA, our estimating equation requires instrumental variables due to the endogeneity/measurement error of the right-side variables. We generally have a common set of instruments for all years. This set consists of unionization, hourly wage,[22] index of intra-industry trade, incentives to freight, nominal and effective subsidies (unrelated to trade protection), and specific components of nominal subsidies for each year subject to availability.[23] More precisely, effective subsidy data is not available for 1988. Labor unions were banned until 1983, and, as one would expect, were effectively nonexistent in 1984. Thus, unionization is not there in the list of instruments for those two years. Incentives to freight are available for only 1988. Finally, the index of intra-industry trade is not available for 1990.

Unionization data (UNION) are for 1994 and used for each year except 1983 and 1984 (when unionization was nonexistent). They are obtained from the Household Labor Force Survey of Turkish State Institute of Statistics (SIS) (http://www.die.gov.tr).

The data on nominal hourly wage rate in each industry are obtained for 1994 from the Employment and Wage Structure Survey of SIS. The nominal wages for other years are constructed from the 1994 data by adjusting it using the inflation rate (based on the GDP deflator), as each year workers' wages are increased based on inflation.

The data on nominal subsidies (and some components), effective subsidies (modified),[24] index of intra-industry trade, and import growth are from Togan (1994).

A. *Construction of the Political Organization Variable*

Because no data were available on trade-related (or other) political contributions, the political organization dummy variable is constructed by other means. Our approach involves two steps: in the first step, membership data for the Turkish Industrialists and Businessmen Association (TUSIAD) are obtained,[25] and an initial determination of organized sectors is made. In the second step, we use discriminant analysis methods to statistically validate the choice made in the first step.[26] We next describe each step in some detail.

After mapping individual members of TUSIAD to their respective sectors, we count the members per sector. Using a cutoff of at least five members, we classify twelve of the 37 sectors as organized.[27] We then augment this list by an additional four sectors with fewer than five members each

[20] It is important to note here that we are covering 86.2% of all imports of 1990.

[21] NRP is the total customs duties plus other charges and expenses related to imports divided by the c.i.f. value of the imports of a particular commodity. ERP is the percentage increase in value added solely due to the presence of import protection measures (both on output and inputs). Finally, NTB is the share of imports "subject to permission" in sectoral imports.

[22] In the Grossman-Helpman (1994) model, the wage rate is exogenous to protection, as it is solely determined by the technology of the Ricardian numeraire sector. As explained later in this section, we use nominal wages that (especially in developing countries) are determined through minimum-wage legislation, which in turn depends on the cost of subsistence.

[23] Except for the case in which a complete system of structural equations is specified, finding appropriate variables to serve as instruments is a difficult problem. Although economists can write new models in which they endogenize variables that were previously treated as exogenous, we strictly follow the Grossman-Helpman model to set our minimum standards in classifying variables as endogenous or exogenous. Thus, as in the Grossman-Helpman model, all policies or incentives other than trade protection are treated as exogenous to the model. In fact, all the variables that we use as instruments are truly exogenous in the context of the Grossman-Helpman model. This is in contrast to the choice of instrumental variables in the existing empirical literature on endogenous trade protection. For example, in Goldberg and Maggi (1999) and Gawande and Bandyopadhyay (2000), the list of exogenous variables includes factor shares, sectoral unemployment, sectoral employment size, output growth, firm scale, and so on that can be considered "somewhat endogenous," even purely in the context of the Grossman-Helpman model.

[24] We construct purely nonimport protection effective subsidy rates that capture the relative increase in value added solely through domestic subsidies or incentives (over the value added in the absence, solely, of these measures) from those presented by Togan (1994), which effectively combine the effective rate of protection with the nonimport protection effective subsidy rate.

[25] We are grateful to Mr. Abdullah Akyuz, the Washington DC representative of TUSIAD, for his generosity in providing us with the list of TUSIAD members.

[26] TUSIAD is a private organization consisting of 470 individual members that hold business positions in a variety of sectors. Large import-competing firms are heavily represented. The organization is very active in Turkish public life, and some of its members are household names. It has representative offices in Washington DC and Brussels and publishes its own newsletter and quarterly economic survey.

[27] The cutoff point has been selected by looking at the frequency tabulation of members per sector.

but whose members are well known for their political and economic clout (based on national newspaper reports).

Because this choice of organized sectors contains elements of subjective judgment, we next examine whether our choice could somehow be statistically validated. We try two alternative methods: discriminant analysis and probit regressions. We see which sectors were ex ante misclassified and calculate the classification error.

A brief summary of the discriminant analysis procedure that we use is as follows. The overall set of all sectors is partitioned into two subsets: subset 1 (organized) and subset 2 (unorganized) based on TUSIAD membership data as explained previously. There is theoretical literature that deals with the (measurable) characteristics that are correlated with whether a sector is organized or unorganized. These measurable characteristics are in most cases the determinants of protection (z and e) and protection itself.[28] In addition, we use other measurable characteristics (that determine the extent of ease or difficulty in organizing): namely four-firm concentration ratios, the data for which have been obtained from the SIS Web site. Therefore, based on the sample means and correlations of all these characteristics of organized and unorganized subsets respectively, we estimate their multivariate normal joint density functions separately for the organized population (of sectors) and for the unorganized population. For our initial classification/partitioning (based on TUSIAD membership data) to be validated, the two joint density functions and their respective estimated parameter vectors should be significantly different (which is what we check in this analysis). To compute the extent of error in our classification, we use a two-stage procedure. Based on the two estimated density functions, we first perform a new ex post classification and compare this new one with our ex ante classification whose percentage error we then calculate.

In general, the discriminant analysis results support our ex ante choice of politically organized sectors, with an average ex post apparent error rate of less than 23%, which is fairly small for a sample size of 37 sectors.[29]

We also used some probit regressions to further scrutinize our classification. The dependent variable was the political organization dummy, and the right-side variables were the import penetration ratio and the import demand elasticity, both purely import-related variables. The variables are jointly significant and have the expected signs (negative for both the import demand elasticity, e, and the import penetration ratio, $1/z$), e being individually significant at the 1% through 5% levels for the last two years and the 5% through 10% levels for the first two years. We then construct an ex post classification by categorizing a sector as organized if the predicted probability of being organized (using the estimated probit regression) is 0.6 or higher. The average percentage error (from misclassification) in this case is around 26.75% and was as low as 24% for 1983 and 1990.

We also do some sensitivity analysis in which we replace in our main tariff regressions the ex ante data on political organization with the ex post classification from our discriminant analysis and our probit regressions. The results are discussed in the section on sensitivity analysis.

V. Results

As explained in section III, we estimate a and α_L, our parameters of interest directly using nonlinear 2SLS and GMM. We present year-specific as well as pooled/panel regressions. We have one set of pooled regressions that includes all years and another one that has all years except 1983. We present results obtained with NRP in considerable detail and then outline the main results with ERP and NTB.

A. *Results with NRP*

Table 1 presents our results with NRP/(1 + NRP) as the dependent variable. In other words, NRP is the empirical measure of t in our equilibrium tariff equation. As can be clearly seen, for all single-year equations, α_L is very tightly estimated, in all cases significant at the 1% through 5% levels, except for 1984. As has been explained in section II, α_L is the proportion of the population that is politically organized. Barring 1984, our estimates of α_L lie in the range of 0.65 to 0.80, implying that 65% to 80% of the population was politically organized. Note that these estimates lie in the economically meaningful range (0, 1), although no such restrictions were placed during estimation. For 1984, we have a very low figure of 0.29, but it is just marginally significant. The 95% confidence interval permits a value of α_L up to 0.76, whereas the corresponding value with a 90% interval is 0.69. All these estimates of α_L are much lower than the 88% and 95% obtained, respectively, by Goldberg and Maggi (1999) and Gawande and Bandyopadhyay (2000) for the United States. Our estimates of a (the weight on welfare relative to contributions) for the single-year equations are fairly precise, although not as tight as those of α_L. These estimates for various years lie in the range 76 through 104. These figures look quite high, but they are comparable to those of Goldberg and Maggi (1999) for the United States and are far lower than those of Gawande and Bandyopadhyay (2000).[30] The high magnitude of the

[28] The expected protection on being organized and the protection on remaining unorganized are important determinants of whether members of a sector decide to get organized. (See Mitra (1999).) Holding all other characteristics constant, there should be a substantial difference between protection received by an organized sector and that received by an unorganized sector with the same set of other characteristics.

[29] It is interesting to note that across years almost all of the organized sectors were correctly classified, so that missclassification occurred almost exclusively in sectors that we characterized as unorganized.

[30] Goldberg and Maggi obtain a value of $\beta = a/(1 + a)$ that equals 0.986, which implies an estimated a equal to 70, whereas Gawande and Bandyopadhyay obtain an estimated a equal to 3,175. McCalman (2000) has estimates in the range of 41 to 47. Note that, for our results, the t-ratio associated with Goldberg and Maggi's β is given by $(1 + a)$ times the

TABLE 1.—ESTIMATION RESULTS FOR NRP/(1 + NRP) EQUATION

	Single Year Results (NL2SLS)					
Year	c	α_L	a	D_{84}	D_{88}	D_{90}
1983	0.36***	0.65***	76.30*	n.a.	n.a.	n.a.
	(0.03)	(0.21)	(42.42)			
1984	0.36***	0.29*	92.43**	n.a.	n.a.	n.a.
	(0.03)	(0.24)	(43.13)			
1988	0.34***	0.67**	63.82*	n.a.	n.a.	n.a.
	(0.03)	(0.28)	(38.96)			
1990	0.24***	0.80***	104.35*	n.a.	n.a.	n.a.
	(0.04)	(0.29)	(58.17)			
	Panel Results					
	c	α_L	a	D_{84}	D_{88}	D_{90}
NL2SLS (DV)	0.36***	0.68**	85.11**	0.03*	−0.02	−0.14***
	(0.04)	(0.29)	(35.54)	(0.03)	(0.03)	(0.03)
NL2SLS (DV)	0.39***	0.67**	80.24**	n.a.	−0.05	−0.16***
(excl. 1983)	(0.04)	(0.34)	(40.94)	n.a.	(0.03)	(0.03)
	c_{83}	α_L	a	c_{84}	c_{88}	c_{90}
NL2SLS (FE-SI)	0.37***	0.76***	96.84***	0.39***	0.35***	0.23***
	(0.03)	(0.20)	(36.47)	(0.02)	(0.02)	(0.03)
NL2SLS (FE-SI)	0.38***	0.61**	85.35**	n.a.	0.34***	0.22***
(excl. 1983)	(0.03)	(0.26)	(35.34)	n.a.	(0.03)	(0.03)
GMM (FE-SI)	0.38***	0.88***	85.95***	0.39***	0.35***	0.24***
	(0.01)	(0.06)	(17.39)	(0.01)	(0.01)	(0.01)
GMM (FE-SI)	0.39***	0.91***	89.00***	n.a.	0.35***	0.24***
(excl. 1983)	(0.02)	(0.25)	(31.23)	n.a.	(0.02)	(0.02)

DV stands for regressions with dummy variables, FE for fixed effects, and SI for regressions with year-specific instruments.
Standard errors in parentheses. * represents absolute value of the t-ratio greater than 1, ** for an absolute value greater than 2, and *** for greater than 2.5.
Overidentifying restrictions accepted at the 5% level for all cases, except 1984 wherein they are accepted at the 4% level (test described in section III).

estimates of a and α_L (in this as well as previous studies) can arise from, among other things, inadequate disaggregation of the data as well as restriction of focus only on trade policy in the Grossman-Helpman model.[31]

The first year in our data set is 1983, which is the last year under dictatorship, following which democracy returned to Turkey. The estimated a and α_L for the year 1983 are 76.3 and 0.65, respectively. The estimated value of a is higher for all other years except 1988. From our single year results, the mean value of the estimated a for the democratic period is around 87 (higher than the estimate for the dictatorship year 1983), and the mean value of α_L for this period is 0.59.[32] Thus the average estimated proportion of the population organized was lower in the democratic regime. This is quite consistent with the predictions of Mitra (1999), which shows that an increase in a leads to a reduction in the equilibrium number of lobbies and thus in the proportion of the population that is organized. This arises from the reduction in the incentives to lobbying and lobby formation (and possibly lobby maintenance) from the reduction in the government's affinity for political contributions as reflected in an increase in a. If we plug in the estimates of a and α_L into $(1 - \alpha_L)/(a + \alpha_L)$, which is the sum of the coefficients of z/e and Iz/e, we see that the values are roughly the same (around 0.45) for both 1983 and for the post-1983 period on the average, implying that, in an organized sector with the same characteristics, z and e would receive the same protection in 1983 as in the democratic period. This is the exit effect of the organized population: as the organized population shrinks in response to an increase in a, the remaining organized population faces less competition, a force on organized sector protection acting in a direction opposite to the direct effect of the reduction in a. However, we do observe lower rates of protection for both organized and unorganized sectors in the democratic period because this was the period in which trade reforms were undertaken. This is because z_i/e_i generally declined from a mean value of 8.55 to a mean value of 6.89. Further, the constant term is estimated to be lower in the post-1983 period.

We now look at the panel regressions. From our dummy-variable regressions with a common set of instruments across years, we estimate the value of a to be 85 for all years pooled, whereas it is 80 for all years pooled except 1983. Both estimates are very tight. Again, the values are higher

t-ratio of our estimated a. Thus, if we were to use β instead of a, we would obtain highly significant results for all cases examined.

[31] See a more detailed discussion in subsection VD.

[32] As can be seen from tables 1 and 3, dropping 1983 from our panel lowers the estimate of a. Two forces drive this result: one comes from the additional variation in the LHS and RHS variables across years (that is, as we move from other years to 1983 or vice versa), and the other comes from the possibility that these relative variations within 1983 might be different from those for all other years combined. It is the latter kind of variations we should exclusively focus on when we compare dictatorship (1983) with democracy (the post-1983 period). Thus, the appropriate way to do this comparison is to compare the single-year equation for 1983 with the panel regressions for all years other than 1983 pooled together.

TABLE 2.—TEST STATISTICS FOR NRP/(1 + NRP) EQUATION

	Single Year Results (NL2SLS)		
$H_0 \rightarrow$	$1/a = 0$	$\alpha_L/(a + \alpha_L) = 0$	$\alpha_L = 0,\ 1/a = 0$
$H_1 \rightarrow$	$1/a > 0$	$\alpha_L/(a + \alpha_L) > 0$	$\alpha_L > 0,\ 1/a > 0$
Year	t-ratio	t-ratio	F-test
1983	1.80	1.22	4.98
1984	2.15	0.92	2.39
1988	1.64	1.53	5.33
1990	1.79	1.31	4.12
	Panel Results		
$H_0 \rightarrow$	$1/a = 0$	$\alpha_L/(a + \alpha_L) = 0$	$\alpha_L = 0,\ 1/a = 0,\ \delta = 0$
$H_1 \rightarrow$	$1/a > 0$	$\alpha_L/(a + \alpha_L) > 0$	$\alpha_L > 0,\ 1/a > 0,\ \delta \neq 0$
Year	t-ratio	t-ratio	F-test
NL2SLS (DV)	2.40	1.60	5.05
NL2SLS (DV) (excl. 1983)	1.96	1.51	10.65
NL2SLS (FE-SI)	2.66	2.41	240.18
NL2SLS (FE-SI) (excl. 1983)	2.41	1.81	184.29
GMM (FE-SI)	4.95	4.70	1053.05
GMM (FE-SI) (excl. 1983)	2.85	3.81	786.57

Table entries are test statistic values.
All alternative hypotheses are one sided, and critical values should be used accordingly.

than for 1983 alone. When we look at the regressions with year-specific fixed-effects (equivalent to dummy variables) and year-specific instruments, a is tightly estimated and varies from 85 to 97, depending on the specification. In any event, these estimates are much higher than for the single-year equation for 1983 alone.

For our dummy-variable regressions with a common set of instruments across years, the estimated value of α_L (again highly significant) is 67% to 68%, roughly the same as that for 1983 alone. When we look at the regressions with year-specific fixed-effects (effectively equivalent to dummy variables) and year-specific instruments, the estimated α_L's are 0.76 and 0.88 with NL2SLS and GMM, respectively, and are significant at the 1% level. When we drop 1983, the estimates are 0.61 and 0.91, respectively, and again extremely tight. The GMM estimates seem to be unreasonably high for Turkey. However, they are close to those obtained in the studies for the United States.

Following estimation, we perform certain tests. As can be seen from table 2, the model is generally significant at the 1% through 5% levels except for the single-year equation for 1984, which is significant at roughly the 10% level.[33] Furthermore, we reject the null hypothesis that $1/a = 0$ (the government is an aggregate welfare maximizer) against the one-sided alternative that $1/a > 0$ at the 1% through 5% levels in all cases. This test here is also roughly equivalent to looking at the significance of the coefficient of z/e, which is $1/(a + \alpha_L)$. As we can see, positive estimates of a and α_L (as we have obtained) produce positive values of $1/(a + \alpha_L)$, which is consistent with the Grossman-Helpman prediction that the coefficient of z/e is positive. Another Grossman-Helpman prediction of course is that the coefficient of Iz/e is negative. This coefficient is $(-\alpha_L/(a + \alpha_L))$, which is negative for positive estimates of a and α_L. We, therefore, test the null hypothesis that $\alpha_L/(a + \alpha_L) = 0$ against the alternative that $\alpha_L/(a + \alpha_L) > 0$. We reject the null at the 1% through 10% levels of significance for all the regressions except those for the years 1983 and 1984, which are significant at the 12% and 18% levels, respectively.

Based on these estimates of a and α_L (and their standard errors), we have also computed the composite coefficients (the coefficients of z/e and Iz/e) of the NRP equation (and their standard errors by the delta method). $-\alpha_L/(a + \alpha_L)$ is the coefficient of z/e, and it (as predicted by the theory) has a negative sign. This coefficient estimate has a small magnitude (less than but close to 0.01 in most cases) due to the very high estimated values of a. Almost all the estimates (single year and panel) are fairly precise (mostly significant at the 1% to 10% levels). The other composite coefficient (of Iz/e), $1/(a + \alpha_L)$, is estimated with a positive sign as predicted by the theory, and it is as significant as the coefficient of z/e. Again, this coefficient estimate has a small magnitude (around 0.01 in most cases), again primarily due to the very high estimated values of a.

B. *Results with ERP and NTB*

Table 3 presents the highlights of our results with ERP/(1 + ERP) and NTB/(1 + NTB) as dependent variables. As the Grossman-Helpman model is about tariffs (NRPs), results with these alternative protection measures should be considered as nothing more than robustness checks (or sensitivity analysis).

With ERPs, α_L is again very tightly estimated (ranging from 0.58 through 0.99 across years and specifications), in all cases significant at the 1% through 5% levels, except for the 1984 single-year equation (which has a marginally significant estimate of 0.24). Our estimates of a are very precise for the fixed-effects regression with year-specific instrument sets. The other regressions for ERP do not yield estimates of a that are as tight, but they still have reasonable significance. All these estimates of a for various years and from the different panel specifications lie in the range 46 through 89.

The estimated a for the year 1983, which is 47.49, is lower than the mean value (of around 54) of single-year estimates of a for the democratic period as well as all other years separately except 1984 (which has almost the same value (45.83)). The value of a is estimated to be 51–89 for all years pooled, and it is 48–77 for all years pooled except 1983, which again is higher than the estimate for 1983 alone.

[33] For our nonlinear estimation, the significance of our model is given by the rejection of the null hypothesis that both $1/a$ and α_L simultaneously equal zero against the alternative that they are both greater than zero. For the panel regressions, there is an additional component $\delta = (\neq)0$ in our null (alternative) hypothesis.

TABLE 3.—ESTIMATION RESULTS

Year	c	α_L	a	D_{84}	D_{88}	D_{90}
ERP/(1 + ERP) equation: Single Year Results (NL2SLS)						
1983	0.41***	0.58**	47.49*	n.a.	n.a.	n.a.
1984	0.41***	0.24*	45.83**	n.a.	n.a.	n.a.
1988	0.43***	0.66**	52.22*	n.a.	n.a.	n.a.
1990	0.37***	0.94***	62.66*	n.a.	n.a.	n.a.
Panel Results						
	c	α_L	a	D_{84}	D_{88}	D_{90}
NL2SLS (DV)	0.44***	0.84**	80.49**	0.03	−0.01	−0.09**
NL2SLS (DV) (excl. 1983)	0.47***	0.81*	76.82*	n.a.	−0.04	0.12***
	c_{83}	α_L	a	c_{84}	c_{88}	c_{90}
NL2SLS (FE-SI)	0.46***	0.99***	80.83*	0.47***	0.45***	0.37***
NL2SLS (FE-SI) (excl. 1983)	0.46***	0.74**	62.04**	n.a.	0.43***	0.35***
GMM (FE-SI)	0.45***	0.92***	51.29***	0.46***	0.45***	0.36***
GMM (FE-SI) (excl. 1983)	0.45***	0.83***	47.94***	n.a.	0.44***	0.34***

NTB/(1 + NTB) equation

Year	c	α_L	a
1984	0.29***	0.85*	66.19*
1988	0.08***	0.89***	47.56*

See notes of table 1.

We only have two years (1984 and 1988) of data for NTBs for which α_L is very tightly estimated at 0.85 and 0.89, respectively. Our estimates of a are not as tight and are 47 and 66, respectively. As explained earlier, consistent with the predictions of Mitra (1999), the year that has a higher a with NTBs is the year with the lower α_L.

Again, in the case of the ERP and NTB regressions, we test the same hypotheses as the ones for NRP. The results are qualitatively very similar. Furthermore, the estimates of the composite coefficients are small in size, have the correct signs, and exhibit a fair amount of precision.

C. Sensitivity Analysis

We experiment with a number of different specifications. Firstly, as explained in section II, to capture the "vote delivering ability" of the agricultural sectors we experiment with an alternative specification as given in equation (5). The additional term is the interaction of the agricultural dummy with z/e, whose estimated coefficient turns out to be statistically insignificant.

In our NTB equation, we also experiment with scaled NTBs. As in Goldberg and Maggi (1999), we use scaling factors of 2 and 3. These scaled-up NTBs provide estimates of α_L greater than 1, and therefore we do not present those results.

In our regressions, we treat both the import demand elasticities and the political organization dummy as endogenous variables. We experiment with regressions that treat both these variables as exogenous, one at a time as well as both together. The results remain qualitatively unchanged. In fact, results are less sensitive to treating elasticities as exogenous than to treating the political organization dummy as exogenous.

We also experiment with some additional exogenous variables (the hourly wage, the degree of unionization, index of intra-industry trade, and the growth rate of imports) thrown into the right side of our estimating equation. All these variables turn out to be individually and jointly insignificant in all our regressions.

The last set of sensitivity analysis checks was performed by reestimating all equations using two different political organization dummy variables: the first was obtained from the ex post classification results of our discriminant analysis, and the second was obtained from the probit regressions of the ex ante membership-based classification on trade-related variables, import penetration ratio, and the import demand elasticity. In general, the results are robust to alternative measures of political organization. In most of these regressions, a and α_L were significantly estimated (at 1%–5% levels), with estimates in the range 55–100 for the former and 0.6–0.9 for the latter.

D. A Broader Interpretation of Results

The estimates of a and α_L in our study as well as in other papers are very high.[34] Certain features are common to our paper and other papers in the literature that drive such results.

[34] If concentration ratios are high and large industries are organized, it is possible for a large proportion of the relevant (to the sample) population to be organized. We are grateful to an anonymous referee for having pointed this out to us.

Firstly, the degree of disaggregation of the data may not be high enough. At the low level of disaggregation, import demand elasticities are low. Additionally, within a sector, certain subcomponent sectors may be organized and others may be unorganized. Treating such whole sectors as organized is equivalent to treating a bigger proportion of the economy as organized. Also, the tariff rates of the subcomponent organized sectors may be much higher than those of the sectors at the more aggregated level. Following the tradition of economic theory, the Grossman-Helpman model focuses on only one type of policy. However, in the real world, the government provides to its favored sectors a wide variety of concessions and services, abstracting from which in empirical analysis may result in overestimates of a and α_L.

VI. Conclusions

The Grossman and Helpman (1994) "Protection for Sale" model theoretically analyzes the determinants of cross-industry protection. They use a "political contributions" approach within a multisectoral, small, open economy setting. Using three-digit, industry-level data from a "genuine" small, open economy (Turkey), we empirically investigate and find support for the fundamental predictions of the Grossman-Helpman model. Our data set for four different years in the period 1983–1990 spans both dictatorial and democratic regimes.

We look at three kinds of protection measures: nominal protection rates, effective protection rates, and NTB coverage ratios. Thus, unlike the two well-known studies for the United States, the set of protection measures we study includes tariffs (nominal protection rates), which is the policy instrument of focus in the Grossman-Helpman model. Further, we perform cross-sectional, single-year regressions as well as panel regressions (with the data pooled across all the available years), thereby checking robustness across years and political regimes.

Our paper also features a few methodological advances. We use classification methods based on discriminant analysis to statistically validate the division of industries into organized and unorganized. We put the estimation problem in its natural, nonlinear form and estimate the structural parameters (and their standard errors) directly. Our pooled equations are estimated both by nonlinear 2SLS and generalized methods of moments (GMM).

As in the previous studies, we find that the government attaches a weight on welfare that is several times the size of its weight on political contributions. Additionally, we find that the weight on welfare relative to contributions was higher in the democratic regime than for dictatorship. We think this result is potentially of importance to researchers in all areas of political economy, and particularly to those studying the relationship between democracy and development.

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APPENDIX A: SUMMARY OF DISCRIMINANT ANALYSIS METHODOLOGY AND RESULTS

Here, we provide a brief summary of the discriminant analysis methodology we follow and some of its results. (See chapter 6 of Anderson (1984) and chapter 11 of Johnson and Wichern (1998) for details.)

Consider the measurement variables mentioned in the text: the protection rate, the inverse import penetration ratio, the (absolute value of the) import elasticities, and the four-firm concentration ratio. Let X denote the $(n \times 4)$ matrix of observations on these variables, and let x_i' denote the ith row of X, $i = 1, 2, \ldots, n$. Using our TUSIAD membership-based ex ante classification variable, n_1 $(=16)$ and n_2 $(=n - n_1 = 21)$ observations (rows) of X come from the organized sector population and unorganized sector population, respectively. Our objective is to try to validate ex post the initial classification using the observations of the measurement variables.

The prior probabilities of classification are estimated from the ex ante separation of sectors as $\pi_1 = n_1/n \approx 0.43$ and $\pi_2 = n_2/n \approx 0.57$. These estimated prior probabilities are taken as the "true" frequencies of organized and unorganized sectors in Turkey. We assume equal cost of missclasification for both organized and unorganized sectors. Let $f_i(x)$ denote the (multivariate) normal density for $i = 1, 2$. If the parameters of the densities are known it can be shown that the region defined by

$$\frac{f_1(x)}{f_2(x)} \geq \frac{\pi_2}{\pi_1}$$

minimizes the expected missclasification cost.

This rule can be applied to the observations of X by substituting estimates for the (true) means and variances of the two normal densities. In most practical applications, it is assumed that the two densities have the same covariance matrix. It can then be shown that the ex post classification rule reduces to "classify x_i to the population of organized sectors if"

$$(\bar{x}_1 - \bar{x}_2)'\hat{\Sigma}^{-1}x_i - \frac{1}{2}(\bar{x}_1 - \bar{x}_2)'\hat{\Sigma}^{-1}(\bar{x}_1 + \bar{x}_2) \geq \ln\left(\frac{\pi_2}{\pi_1}\right)$$

where $\bar{x}_i$ denotes the sample mean of the observations belonging to the ith population and $\hat{\Sigma}$ denotes the estimator of the common covariance matrix, the estimates being derived using our ex ante classification. Thus, from the preceding classification rule, a requirement for successful classification is a substantial difference between the means of the two normal distributions, thereby requiring a test for their equality. The p-values (averaging 0.117) from the tests of equality of means show evidence that the means are not the same in the populations of organized and unorganized sectors.

After the ex post classification, we can compute the posterior probabilities, $\hat{\pi}_1 = \hat{n}_1/n$ and $\hat{\pi}_2 = \hat{n}_2/n$, and compare them to the prior probabilities from the ex ante classification (where the "hat" represents ex post values arrived at using the preceding rule). One can also compute the ex post apparent error rate (AER) of missclassification given by the number of missclassified observations as a proportion of the sample size. The posterior probabilities (with an average value of 0.45 for the organized sectors compared to a prior of 0.43) track the (estimated) prior probabilities quite well. There is a tendency to assign unorganized sectors to organized ones, which inflates the posterior probabilities of the organized sectors. The ex post apparent error rates are relatively low, averaging 0.227 (0.21, 0, 17, 0.29, and 0.24 for the years 1983, 1984, 1988, and 1990, respectively). Because the samples used in the classification are quite small, we feel that the above results support our ex ante classification.