Universität Trier

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Research Papers in Economics
No. 1/11

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February 11, 2010

Abstract

We provide a theory of trade policy determination that incorporates the protectionist bias inherent in majoritarian systems, suggested by Grossman and Helpman (2005). The prediction that emerges is that in majoritarian systems, the majority party favors industries located disproportionately in majority districts. We test this prediction using U.S. data on tariffs, Congressional campaign contributions, and industry location in districts represented by the majority party over the period 1989–97. We find evidence of a significant majority bias in trade policy: the benefit to being represented by the majority party appears at least as large in magnitude as the benefit to lobbying.

Keywords: trade policy, majority bias, political economy.

JEL: F13.

"If they [politicians] are successful, they claim, as a matter of right, the advantages of success. They see nothing wrong in the rule, that to the victor belong the spoils of the enemy."
(New York Senator William L. Marcy, referring to the victory of the Jackson Democrats in the election of 1828, in the U.S. Senate, January 25, 1832)[1]

1 Introduction

In majoritarian electoral systems such as the U.S., politics is particularly grounded in local interests (Milesi-Ferretti, Perotti, and Rostagno (2002)). Grossman and Helpman (2005) (henceforth, G-H) argue that the pattern of protectionism is influenced by the industry structure of the majority-party legislators' home districts. We provide a simple theory of the determination of trade policy that merges G-H's (2005) view of majoritarian system trade politics with the lobbying approach pioneered by Grossman and Helpman (1994).

In our model, the majority delegates consider national welfare, but attach additional weight to the welfare of their own home districts, simultaneously as organized industry lobby groups offer the majority delegation campaign contributions in return for favorable trade policies. Our approach facilitates a comparison of the influence of electoral rules (majoritarian system) relative to the impact of lobbying. The novel predictions are: (i) if an industry is relatively concentrated in majority districts, it receives positive protection (even without an organized lobby); (ii) on the other hand, industries located primarily in minority districts suffer from lower, possibly negative, trade protection.

We test these predictions using a newly assembled unique data set of tariffs and Congressional campaign contributions for 332 U.S. manufacturing industries over the

1989–97 period. This time period encompasses both major trade policy decisions in the U.S., such as the Uruguay Round and NAFTA negotiations, and a change in the majority party from Democrats to Republicans in both Houses of Congress in 1995.[2] We use industry-level employment by district to construct the extent to which the majority party represents each industry, and use political contributions to the majority party to classify industries as organized (i.e., lobbying) or unorganized. All explanatory variables are instrumented similarly to Goldberg and Maggi (1999) (G-M), Gawande and Bandyopadhyay (2000) (G-B), Eicher and Osang (2002), and Matschke and Sherlund (2006).[3]

Using a two-step GMM approach, we find both year-by-year and panel data estimates consistent with the model's predictions. The results are robust across three different lobby group organization classification methods, following the approaches of G-M, G-B, and Ederington and Minier (2008) (E-M).

There have been some recent studies that also explicitly consider the impact of the electoral system on trade policy. Baldwin and Magee (2000) find that the structure of industry employment in the home districts of members of the House of Representatives affected voting behavior on the General Agreement on Tariffs and Trade (GATT) and on Most Favored Nation (MFN) status for China in 1993-94 to some degree. In particular, higher employment in the textiles industry reduced the likelihood of a legislator voting for the GATT and MFN bills. Although there was no significant impact on most legislators' voting behavior, Baldwin and Magee argue that excluding industry characteristics in studies of U.S. trade policy voting yields omitted variable bias. Conconi, Facchini, and Zanardi (2009) use county-level industry data to study the voting behavior of members of the U.S. House and Senate on the decision to grant Fast Track Authority (FTA) to the President on trade agreements. A member of Congress is more likely to vote in support of FTA, the more export-oriented is her

constituency relative to the national average. Concerning the effect of a majoritarian electoral system, Willmann (2005) studies a theoretical model of a majoritarian system with all regional representatives in the national legislature affecting trade policy determination equally. Each industry sector is concentrated in one electoral district only. Willmann finds that strategic delegation induces each district to elect a legislator who is more protectionist than the median voter, resulting in positive tariff levels. In related work, Roelfsema (2006) shows that a majoritarian electoral system results in a higher level of trade protection than does a proportional electoral system; Evans (2009) finds evidence that countries with majoritarian systems have, ceteris paribus, higher average tariff levels. To our knowledge, however, our study is the first to empirically evaluate the impact of majoritarian systems on trade policy outcomes at the industry level as suggested by Grossman and Helpman (2005).

Our results indicate that electoral rules are an important determinant of trade policy; in particular, the benefit to being located in districts represented by the majority party appears to be larger in magnitude than the well-established benefit to lobbying. Our paper complements, therefore, Lizzeri and Persico (2001), Milesi-Ferretti et al. (2002), and Persson and Tabellini (2004), for example, who find policy effects of electoral rules on public spending.[4]

The paper is organized as follows. Section 2 sets up the theoretical model and derives the predictions. Section 3 describes the data. Section 4 outlines the empirical approach and reports the results, and Section 5 concludes. Appendix A provides further details on variable construction, and in Appendix B, we present additional results when majority representation is calculated at the state (rather than Congressional district) level.

2 Model

A small open economy has individuals living in N geographically separate political districts indexed by j. The population is normalized to unity. Each individual i in this economy consumes $n+1$ goods and has quasi-linear preferences given by $x_0^i + \sum_{g=1}^{n} u(x_g^i)$, where x_0^i represents i's consumption of the numeraire good 0 and $u(.)$ is a differentiable and strictly concave function of consumption x_g^i of good g, $g \in \{1, 2, ..., n\}$. Good 0 has world and domestic price equal to unity, while other goods g have world and domestic prices p_g^w and p_g, respectively. Sector g is protected by a specific import tariff or export subsidy; i.e. $p_g = p_g^w + t_g$.

Individual i living in district j spends $\sum_{g=1}^{n} p_g d_g(p_g)$ on non-numeraire goods, where demand for good g is given by $d_g(p_g) = [u'(x_g)]^{-1}$; we drop individual-specific superscripts (consumption quantities of all non-numeraire goods and associated consumer surplus are equal across individuals, provided that all individuals have sufficient income). The remaining budget share is spent on numeraire good 0; this amount is assumed strictly positive.

Good zero is produced from labor only with constant returns to scale and an input-output coefficient equal to unity; assuming positive production, the wage rate equals one. Good $g \neq 0$ requires labor and a sector-specific input. With a fixed wage rate, the aggregate factor reward in sector g, $\pi_g(p_g)$, depends on p_g only. Each individual receives wage income. The consumer surplus derived from good g consumption equals $s_g(p_g) = u[d_g(p_g)] - p_g d_g(p_g)$. Tariff revenue collected in sector g equals $r_g(p_g) = (p_g - p_g^w) m_g(p_g)$, where $m_g(p_g) = d_g(p_g) - X_g(p_g)$ is the net import demand function and $X_g = \pi'(p_g)$ is the domestic supply of good g by Hotelling's Lemma.

Individuals may own sector-specific input factors in at most one sector g. In some or all of the n sectors, denoted by L, the factor owners organize national lobby groups incorporating capital owners across districts. In organized sectors, sector g's lobby

seeks to influence trade policy by offering campaign contribution schedules $C_g(p_g)$ to the majority legislative delegation (Grossman and Helpman (1994)). With highly concentrated ownership, factor owners value only factor reward. Thus, the gross welfare of the sector g lobby equals

$$W_g(p_g) = \pi_g(p_g). \tag{1}$$

Contrary to the original protection-for-sale model (Grossman and Helpman (1994)), industry g lobby welfare thus only depends on p_g, rather than on the entire price vector.

Denote the district j population share by β_j and the share of industry g capital located in district j by α_{jg}. Then, the aggregate income level of district j equals

$$Y_j = \beta_j + \sum_{g=1}^{n} \alpha_{jg}\pi_g + \beta_j \sum_{g=1}^{n} r_g(p_g), \tag{2}$$

where the RHS terms equal labor income, capital income, and net transfer income, respectively. Adding consumer surplus to Eqn. (2) yields district j residents' aggregate social welfare level,

$$W_j^A = \beta_j + \sum_{g=1}^{n} \alpha_{jg}\pi_g + \beta_j \sum_{g=1}^{n} [r_g(p_g) + s_g(p_g)]. \tag{3}$$

Each district j is represented by a single legislator who is affiliated with either the majority or minority party. Majority delegation legislators value overall citizen welfare, but they put an additional weight ζ on welfare in their home districts. This assumption receives empirical support from the literature on the relationship between distribution of public funds and party control. Levitt and Snyder (1995) report that when the Democrats had a majority in Congress, federal spending in an area increased with its Democratic vote. Ansolabehere and Snyder (2006) find that areas providing the largest vote share to the incumbent party receive the highest shares of state transfers to local government. Moreover, the distribution of funds is redirected

towards the new governing party's core supporters as a result of a change in the state government.[5] As one example of this in practice, Joanis (2009) shows that in Quebec, the geographic allocation of spending is highly dependent on districts' party loyalties.

The majority party is represented by at least $(N/2+1)$ legislators, where N is the total number of districts; the set of majority districts is denoted by K. Majority-party representatives may compensate each other with political side payments or inter-temporal trades; they maximize their joint welfare,

$$W^M = \phi \sum_{g \in L} C_g(p_g) + \zeta \sum_{j \in K} W_j^A + \sum_{j=1}^{N} W_j^A, \tag{4}$$

where the weight $\phi > 0$ is the majority legislators' weight on contributions relative to general welfare and $\zeta > 0$ is the additional weight legislators place on welfare in the majority districts.[6]

The equilibrium trade policy is determined as the outcome of a two-stage, non-cooperative game. In stage one, each organized lobby $g \in L$ simultaneously and non-cooperatively offers the legislative majority a contribution schedule $C_g(p_g)$, taking the other lobbies' strategies as given. We assume, consistent with Grossman and Helpman (1994), that the contribution schedules are differentiable in product prices. In stage two, the legislative majority selects its favored trade policy and collects the associated contribution from each organized lobby; the lobbies are assumed not to renege on their promises in this stage. As described in Grossman and Helpman (1994), the equilibrium trade policy maximizes the sum of policy-maker and lobby welfare. Moreover, the contribution level of lobby g can be easily determined: Since the lobby designs the contribution schedule and the policy-maker can only accept or reject, the lobby of industry g can extract all surplus from the contributions game, leaving the legislature with only the welfare that it would obtain without lobby presence in industry g. We write the equilibrium trade policy for sector g with lobby presence as $t_g^*(\delta_g = 1)$ and the equilibrium trade policy without lobby presence in sector g

as $t_g^*(\delta_g = 0)$, where "*" denotes an equilibrium value and where δ_g is an indicator variable equal to 1 if firms in sector g are organized, and zero otherwise. Lobby g's contributions then equal

$$C_g = \frac{\zeta}{\phi}\left[\sum_{j\in K}[W_j^A(t_g^*(\delta_g=0)) - W_j^A(t_g^*(\delta_g=1))]\right] + \frac{1}{\phi}\left[\sum_{j=1}^{N}[W_j^A(t_g^*(\delta_g=0)) - W_j^A(t_g^*(\delta_g=1))]\right], \quad (5)$$

i.e., contributions just compensate the legislature for the welfare loss due to the lobby distortion.

We now derive the equilibrium trade policy. When maximizing (4) by choice of t_g, the equilibrium characterization equals

$$t_g m_g' + \phi\delta_g X_g(p_g^*) + \zeta\left[X_g\sum_{j\in K}(\alpha_{jg}-\beta_j) + \sum_{j\in K}\beta_j t_g^* m_g'(p_g^*)\right] = 0, \forall g, \quad (6)$$

which, noting that $m_g' < 0$, yields, for all g,

$$t_g^* = \frac{\phi}{1+\zeta\sum_{j\in K}\beta_j}\frac{\delta_g X_g(p_g^*)}{|m_g'(p_g^*)|} + \frac{\zeta}{1+\zeta\sum_{j\in K}\beta_j}\sum_{j\in K}[\alpha_{jg}-\beta_j]\frac{X_g(p_g^*)}{|m_g'(p_g^*)|}. \quad (7)$$

Tariff protection is a function of industry size (X_g) and the absolute slope of the import demand function ($|m_g'|$). Sector g unambiguously receives a positive level of protection if the majority party districts' share of sector g capital is greater than the population share of these districts, $\sum_{j\in K}\alpha_{jg} > \sum_{j\in K}\beta_j$, even if the industry does not lobby. Moreover, if the sector lobby is organized ($\delta_g = 1$), sector g receives positive protection, even if these shares are exactly equal. Finally, even if the majority districts' share of sector g capital is smaller than their population share, sector g may receive positive protection due to lobbying. This occurs only if the majority legislation values campaign contributions sufficiently highly relative to social and majority district welfare (high ϕ). We also see that holding the ratio of output to import demand slope $X_g/|m_g'|$ fixed, the difference between $t_g^*(\delta_g = 1)$ and $t_g^*(\delta_g = 0)$

(that is, the tariff difference for industries with and without active lobbies) increases with higher contributions weight ϕ, lower additional weight ζ on majority district welfare, and lower majority-district population share $\sum_{j \in K} \beta_j$.[7]

3 Data

To estimate the model and test its predictions, we use a panel of U.S. manufacturing industries covering a large part of the 1990s. More precisely, we predict tariff levels in 1993, 1995, and 1997 as a function of industry lobbying and representation by the majority party in 1991, 1993, and 1995. Since this implies a substantial update over much of the previous empirical G-H literature, our results are not directly comparable. Our study is also, to our knowledge, the first contribution to the empirical G-H literature that uses panel data to evaluate the determinants of U.S. trade policy. Summary statistics for our data by year can be found in Table 4 in Appendix A.1.

Our measure of trade protection is tariffs, following the theory. Our tariff and import data come from Schott's (2008) trade database. Data on other industry characteristics are primarily from the Bartelsman, Becker, and Gray (2000) NBER productivity database. The import demand elasticity measures come from Kee, Nicita and Olarreaga (2008). Import demand elasticities are estimated by industry at the 6-digit HS level, which we concord to SIC4 based on import weights (see Appendix A.3).

We use contributions to majority candidates (Democrats for the years 1991 and 1993; Republicans for the year 1995) in both House campaigns during the 1989–90, 1991–92, and 1993–94 electoral cycles to classify industries as organized in each time period. We experiment with alternative classification methods (described further in Appendix A.4) based on G-M, G-B, and E-M. In the G-M classification, we identify a break in the distribution of the contributions data and consider industries with

contribution levels above the breakpoint organized. As robustness checks, we also use the G-B and E-M lobby classifications. For the G-B classification, we regress contributions data on 2-digit SIC industry dummy variables interacted with bilateral import penetration data for trading partners and interpret the coefficient estimates on these interaction terms; a positive relationship between contributions and import penetration ratio within a 2-digit industry is considered evidence of lobbying for the entire group of 4-digit industries in that 2-digit industry. E-M criticize the common practice of considering some industries that make positive contributions as unorganized and simply classify all industries with positive contributions as organized. In practice, this means that almost all industries are considered organized. Summary statistics for the lobby variables can be found in Table 4. From the table, one can see that the lobby variables do not vary much over time. Indeed, when calculating correlations (not reported here), the lobby measures are positively correlated with each other for any given year and across time, so lobbying activity appears quite stable.

In accordance with theory, we use district-level data to construct the variables measuring the extent to which an industry is represented by the majority party. The exact procedure that we use to calculate $(\alpha_{jg} - \beta_j)$ is laid out in Appendix A.2. The variable $(\alpha_{jg} - \beta_j)$ measures the extent to which an industry is over-represented (relative to employment) in a district. To construct the industry-level measure of representation by the majority party, the $(\alpha_{jg} - \beta_j)$ are summed over majority Congressional districts (Democrats in the 102nd and 103rd Congresses; Republicans in the 104th).

In Appendix B, we also discuss our results when majority representation is measured at the state, rather than the district, level. Using a state-level measure of majority representation allows us to include data on the U.S. Senate as well as the House of Representatives (Senators, of course, are elected statewide), and considering

the state level may also reflect some aspects of the way party politics work in practice, as we discuss in Appendix B. For the main body of the text, however, we base our results on districts as the regional unit since contributions to House candidates can be tied more directly to employment in their Congressional district.

4 Empirical Approach

The econometric model is derived from equation (7), which we rewrite in terms of observables:

$$t_g^* \tilde{m}_g' = \frac{\phi}{1 + \zeta \sum_{j \in K} \beta_j} \delta_g \tilde{X}_g + \frac{\zeta}{1 + \zeta \sum_{j \in K} \beta_j} \sum_{j \in K} z_g \tilde{X}_g, \tag{8}$$

where $\tilde{X}_g$ is the value of industry shipments and $t_g^* \tilde{m}_g'$ can be calculated by noting that $t_g^* \tilde{m}_g' = -t_g^* m_g' p_g^* = \tilde{t}_g^* e_g p_g^* m_g / (1 + \tilde{t}_g^*)$, where $\tilde{t}_g^*$ is the ad valorem tariff rate in the industry, $p_g^* m_g$ is the value of imports, and e_g is the absolute value of the price elasticity of import demand. Finally, $z_g = \sum_{j \in K} (\alpha_{jg} - \beta_j)$ is our measure of the extent to which the industry is located in majority districts, discussed in the previous section. Thus, our estimating equation becomes:

$$t_g^* \tilde{m}_g' = \gamma_0 + \gamma_1 \delta_g \tilde{X}_g + \gamma_2 z_g \tilde{X}_g + \epsilon_g \tag{9}$$

where, according to our derivations, $\gamma_1 > 0$ and $\gamma_2 > 0$.[8] Thus, our testable prediction is that the marginal impact of industry size on tariff protection should be conditional on the proportion of the industry located in majority districts. Intuitively, greater industry protection in majority districts is valued more highly in the government's welfare function, so the marginal impact of industry production on tariff protection is increasing in the majority representation of the industry. We can also estimate the ratio of the weight on majority district welfare compared to campaign contributions by forming the ratio γ_2/γ_1. Finally, with the knowledge of the population share in majority districts $\sum_{j \in K} \beta_j$, we can identify ζ and ϕ.

A standard complication that arises in G-H estimation like this is that the value of shipments (and potentially also whether an industry is organized) may be endogenous, requiring instrumental variables estimation.[9] To address potential endogeneity, we estimate equation (9) using two-step optimal GMM. The instruments for our explanatory variables $\delta_g \tilde{X}_g$ and $z_g \tilde{X}_g$ are comparable to those in G-M, G-B, Matschke and Sherlund (2006), and Matschke (2008): physical capital's share of output, industry employment, and, in addition, the majority variable z_g to improve the performance of instruments for the equation explaining $z_g \tilde{X}_g$. We carefully monitor the appropriateness of instruments, using first-stage F-statistics to check for possible weakness of instruments and calculating Hansen's J-statistic to evaluate the validity of instruments.

We begin the analysis by estimating our tariff equation year-by-year. Estimating equation (9) requires constructing measures for both the extent of majority representation of an industry, z_g, and the organization, δ_g, of the industry. A general description of these variables was given in Section 3; additional details are provided in Appendix A. We use the approaches by G-M, G-B, and E-M to assign a value of zero or one to the lobby indicator.

Table 1: Year-by-Year Estimation Results

Year:	1993	1995	1997	1993	1995	1997	1993	1995	1997
lobby variable:		G-M			G-B			E-M	
$\hat{\gamma}_1$	.0088***	.0086***	.0069***	.0091***	.0087***	.0070***	.0083***	.0076***	.0062***
	(.0010)	(.0008)	(.0009)	(.0017)	(.0019)	(.0016)	(.0011)	(.0011)	(.0010)
$\hat{\gamma}_2$	.0584***	.0648***	−.0464***	.0828***	.0751***	−.0578***	.0593***	.0661***	−.0479***
	(.0133)	(.0147)	(.0107)	(.0254)	(.0206)	(.0168)	(.0138)	(.0152)	(.0112)
$\hat{\gamma}_0$	20.44***	34.23***	42.38***	29.69***	43.05***	53.43***	19.15***	33.95***	41.83***
	(6.26)	(8.00)	(11.07)	(10.10)	(12.89)	(15.74)	(7.36)	(9.52)	(12.21)
$\hat{\phi}$	.0091***	.0089***	.0067***	.0096***	.0091***	.0068***	.0086***	.0080***	.0060***
	(.0010)	(.0008)	(.0009)	(.0017)	(.0020)	(.0016)	(.0012)	(.0011)	(.0010)
$\hat{\zeta}$	.0605***	.0674***	−.0453***	.0872***	.0786***	−.0561***	.0615***	.0688***	−.0467***
	(.0142)	(.0159)	(.0102)	(.0282)	(.0226)	(.0158)	(.0149)	(.0164)	(.0106)
number of observations	319	314	316	328	323	325	319	314	316
first-stage F-stat lobby	16.66	14.19	28.85	10.20	7.49	15.59	33.06	37.41	57.90
first-stage F-stat majority	19.42	23.22	12.45	19.30	23.40	12.71	19.42	23.22	12.45
Hansen's J-stat p-value	.6903	.6897	.6915	0.6866	.6291	.6818	.2392	.2544	.2330

Years identify the year in which tariffs (the dependent variable) are measured. Standard errors (in parentheses) for structural parameters were calculated using the delta method. *, **, and *** indicate significance at 10%, 5% and 1% levels. Instruments are industry employment, industry capital, and the majority measure z_g.

Throughout Table 1, the instruments perform well: F-statistics are typically high (above 10) in the first-stage regressions, with only one exception, and the J-statistic p-values do not indicate any problems with instrument validity. Using the reduced-form estimates together with our information on the population percentage $\sum_{j \in K} \beta_j$ living in majority districts (see Table 4), we can also recover the structural parameter estimates $\hat{\phi}$ and $\hat{\zeta}$.

Our main interest lies in the point estimates for majority influence ζ. The results for the majority variable differ significantly by year in an intuitively appealing way. In the tariff equations for 1993 and 1995, the estimates are statistically significant even at the 1% level of significance. For 1993, the point estimates $\hat{\zeta}$ for the additional weight on majority welfare range from 6.05% to 8.72%; for 1995, the values are closer together and lie between 6.74% and 7.86%. For 1997, in contrast, the structural estimate $\hat{\zeta}$ becomes negative and smaller in magnitude, and remains statistically significant. This finding seems to contradict our theory, but can be readily explained. It is the result of the switch in majority party that took place with the 104th Congress: The 1997 tariffs are presumably the first in our sample determined by the (new) Republican majority that took office in 1995. However, the 1997 tariffs have a correlation of about 97.77% with the 1995 values, showing that the time lag we have assumed for the determination of tariffs may still be too short. In contrast, the majority variable z_g has a negative correlation between its 1995 and 1997 tariff equation values of -60.62%. Accordingly, the coefficient estimates on the majority variable turn negative. If we rerun the estimation using the 1995 majority variable to explain the 1997 dependent variable (results not reported here), the majority variable coefficients are once again positive and significant.

The point estimates for ϕ (the weight on contributions in the governmental objective function when domestic welfare receives a weight of 1 or 100%) are very stable

across time, quite in contrast to our results for ζ. They show a slight downward trend from 1993 to 1997 and lie in the overall range from .60% to .96%. The estimates for ϕ appear much smaller than the estimates for ζ, but are always significant at at least the 1% significance level. The low weights on contributions in the governmental welfare function confirm similar findings in the empirical GH literature that the weight on contributions, while statistically significant, is of a very small magnitude.

Next, we combine the data into a panel, controlling for industry fixed effects by demeaning with the time average of an industry and for time fixed effects by allowing the constant term to vary by year. The results for the combined years 1993 and 1995 are presented in Table 2; we discuss the results including 1997 at the end of this section. In the panel specification, we do not allow the coefficients on the lobby and majority variable to change over time, so our coefficient estimates are time averages. Both the lobby and the majority variable coefficients are precisely estimated and positive and statistically significant at the 1% level of significance. There is no weak instrument problem in any of the three specifications, and the J-test of instrument validity does not reveal any problems. As was already evident in the year-by-year estimations, the majority welfare weight estimate $\hat{\zeta}$ exceeds the contributions weight estimate $\hat{\phi}$. In Table 2, the estimates $\hat{\zeta}$ lie between 6.18% and 7.44%, whereas the estimates $\hat{\phi}$ are in the range from 0.79% to 0.89%.

While the point estimates for the majority welfare weight ζ in 1993 and 1995 are much higher than the point estimates for the contributions weight ϕ, it is not clear whether the effect of majority influence on trade policy is really this much larger in the data, because the lobby variable indicator ranges between 0 and 1 with a mean of over 0.7, whereas the majority variable $z_g = \sum_{j \in K} (\alpha_{jg} - \beta_j)$ in the data lies between a minimum of -0.35 and a maximum of 0.46 and is quite centered around zero (see Table 4). To obtain a better understanding of how lobbying influence compares to

Table 2: Panel Estimation Results, 1993 and 1995

Lobby:	G-M	G-B	E-M
$\hat{\gamma}_1$	.0086***	.0085***	.0076***
	(.0006)	(.0013)	(.0008)
$\hat{\gamma}_2$	.0595***	.0712***	.0620***
	(.0078)	(.0136)	(.0086)
$\hat{\gamma}_{1993}$	66.97***	72.76***	70.59***
	(5.68)	(6.90)	(5.79)
$\hat{\gamma}_{1995}$	79.79***	84.74***	82.91***
	(6.87)	(8.35)	(6.91)
implied structural parameters			
$\hat{\phi}$	.0089***	.0089***	.0079***
	(.0007)	(.0014)	(.0009)
$\hat{\zeta}$	.0618***	.0744***	.0644***
	(.0084)	(.0148)	(.0092)
number of observations	633	651	633
first-stage F-stat lobby	20.44	21.98	58.68
first-stage F-stat majority	72.05	74.50	72.05
Hansen's J-stat p-value	.8608	.7737	.3112

Standard errors (in parentheses) for structural parameters were calculated using the delta method. *, **, and *** indicate significance at 10%, 5% and 1% levels. Instruments are industry employment, majority variable z_g, and z_g multiplied with industry capital.

Table 3: Predicted Tariff Changes Due to Changes in Lobby and Majority Variable

specification:	G-M	G-B	E-M
lobby min to max	2.7909	2.7766	2.4637
majority min to max	13.6915	16.4701	14.4044
ratio maj to lobby	4.9057	5.9318	5.8467
lobby mean to mean + 1 std.	1.2440	1.2860	.2005
majority mean to mean +1 std.	1.8031	2.1689	1.8970
ratio maj to lobby	1.4495	1.6866	–

The predicted ad valorem tariff changes are given in percentage points and were calculated using the sample averages for 1993 and 1995 as reported in Table 4. The standard deviation for lobby and majority variable are calculated as weighted average over time rather than over the entire two-year sample in order to allow the reader to reconstruct the results using the provided data.

majority bias with respect to trade policy, we use the two-year panel results reported in Table 2 and conduct two comparative statics exercises.

In the first exercise, we take a fictitious industry with average characteristics and ask how the predicted ad valorem tariff would change if it went from a politically inactive industry to a lobbying industry. Then we calculate a similar tariff change when the industry characteristics change from being an industry least concentrated in majority districts to one with the highest concentration in majority districts in the combined 1993 and 1995 data. The results of this exercise are recorded in the first part of Table 3. With percentage point estimates of the tariff change in the range of 2.46% to 2.79%, the lobby effect on tariffs is actually quite large given the overall low level of tariffs. However, when switching the majority variable from its minimum (-0.33) to its maximum (0.36), holding the lobby variable at its mean, the majority effect is several times larger, by a factor of between 4.91 and 5.93.

We also calculate the effects of less extreme changes, in particular the induced tariff change (reported in the lower part of Table 3) when – starting from the lobby

or majority variable mean – we increase the respective variable by one standard deviation.[10] In this comparative static exercise, we see that the induced tariff changes are much more similar, but still, the induced variation from the majority variable change is between 1.45 to 1.69 times higher than the induced variation from the lobby variable change. We thus conclude that in the early years of the sample (with Democratic majority), the majority effect on trade policy indeed appears larger than the lobby effect.

If we consider the full three-year sample, not surprisingly, the statistical significance of the majority variable disappears.[11] Yet overall, we view the panel results as confirmation that majority bias on trade policy is detectable in our sample, in accordance with our prediction that industries located primarily in majority districts are favored in the political process. The implementation lag, however, seems quite long and leads to a rejection of the majority bias hypothesis for the tariff data of 1997.[12]

5 Conclusion

In this paper, we incorporated majority bias, as suggested by Grossman and Helpman (2005), into a standard Protection-for-Sale model. Trade policy is determined by the majority legislative delegation, which cares about domestic overall welfare, but even more so about majority-district welfare, as well as campaign contributions. In this case, in addition to the industry's lobbying influence, the extent to which the industry is concentrated in majority districts matters for trade policy. If an industry's capital share in majority districts exceeds the districts' population share, it will receive positive protection even if it does not have an active lobby. Industries that are only weakly represented in majority districts receive lower trade protection.

We test these predictions using a newly assembled data set of U.S. manufacturing

industries containing tariff levels in 1993, 1995, and 1997 and attempt to explain them by industry variables, in particular lobby and majority variables, for 1991, 1993, and 1995. The results are supportive of the theoretical model in the sense that both the coefficients for the standard Protection-for-Sale lobbying variable and the majority bias variable are statistically significant and of the theoretically predicted signs for the first two years of our three-year sample, when the Democratic majority in Congress had been in power for a considerable amount of time.[13] We conclude that the structure of the legislative decision-making process — in addition to lobbying influence — matters for the determination of trade policy.

A Data Construction

A.1 Data Summary

Table 4 gives a summary of our data sample. Here, we excluded all observations for which any reported variable was missing, i.e. the presented data are "balanced" by year. Depending on the exact specification, the number of observations to estimate a certain equation may thus be somewhat higher.

Table 4: Data Summary

tariff equation 1993 (319 obs.)	Mean	Std. Dev.	Min.	Max.
tariff	.0537	.0363	0	.0258
import demand elasticity	-1.5203	1.5561	-18.5885	-.1597
imports (in million $)	1302.238	4193.703	.2097	62314.22
shipments (in million $)	6414.5	11307.85	158.6	133861.2
employment (in 1000)	38.0198	51.5851	.5	400.1
capital (in million $)	2809.717	5512.998	31	54488
E-M lobby	.9937	.0791	0	1
G-M lobby	.7649	.4247	0	1
G-B lobby	.7053	.4566	0	1
$z_g = \sum_{j \in K} (\alpha_{jg} - \beta_j)$	-.0001	.0872	-.2722	.3623
$\sum_{j \in K} \beta_j$	.6044	–	–	–
tariff equation 1995 (314 obs.)	Mean	Std. Dev.	Min.	Max.
tariff	.0500	.0501	0.0001	.7001
import demand elasticity	-1.5204	1.5563	-18.5885	-.1597
imports (in million $)	1732.578	5549.547	2.4935	76486
shipments (in million $)	7239.425	13463.3	24.9	167825.8
employment (in 1000)	38.8758	54.9802	.1	443.9
capital (in million $)	2912.575	5686.193	28.7	52533.2
E-M lobby	.9936	0.0797	0	1
G-M lobby	.7293	.4450	0	1
G-B lobby	.7229	.4483	0	1
$z_g = \sum_{j \in K} (\alpha_{jg} - \beta_j)$	-.0096	.0944	-.3302	.3105
$\sum_{j \in K} \beta_j$	.5962	–	–	–

Table continues on next page.

Table 4, continued

tariff equation 1997 (316 obs.)	Mean	Std. Dev.	Min.	Max.
tariff	.0413	.0506	0	.7000
import demand elasticity	-1.5268	1.5615	-18.5885	-.1597
imports (in million $)	1988.113	6181.027	4.1639	87549.91
shipments (in million $)	8386.974	16136.01	0	2012
employment (in 1000)	40.1535	58.8501	0	500.9
capital (in million $)	2998.715	5829.197	26.5	49801.7
E-M lobby	1	0	1	1
G-M lobby	.7215	.4490	0	1
G-B lobby	.7120	.4535	0	1
$z_g = \sum_{j \in K} (\alpha_{jg} - \beta_j)$	.0198	.1012	-.3548	.4638
$\sum_{j \in K} \beta_j$	.5362	–	–	–

A.2 Majority Representation Variable

We constructed a variable measuring the extent of an industry's majority representation as follows. County Business Patterns gives, for each county, employment by 4-digit SIC. Since many observations are censored, it also gives the number of establishments in various size classes. These data are also available at the national level, with fewer censored observations. Following Busch and Reinhardt (1999), for each industry, we compute the mean establishment size at the national level for each size class. Then, for each county, we estimate total employment by industry using the national industry averages for each size class and the number of firms in each size class by county. Following Busch and Reinhardt, we use the imputed data even when the actual data are not censored. The percentage of industry employment in each county is estimated using the sum across counties as the denominator

(so that the percentages sum to one for each industry). We then concord from the county level to the Congressional district level using the MABLE/Geocorr90 Geographical Correspondence Engine provided by the University of Missouri Data Center (http://mcdc2.missouri.edu/websas/geocorr90.shtml). Congressional districts were redistricted between the 102nd and 103rd Congresses (but districts for the 104th are identical to those for the 103rd).

To calculate $(\alpha_{jg} - \beta_j)$, we use the difference between the percentage of an industry's employment located in a given district and the population share of that district (from the 1990 Census). This gives a measure of the extent to which an industry is over-represented (as measured by employment) in a district. To construct the industry-level measure of representation by the majority party, the $(\alpha_{jg} - \beta_j)$ are summed over majority districts (Democrats in the 102nd and 103rd Congresses; Republicans in the 104th). This is our measure of $(\alpha_{jg} - \beta_j)$. We use majority representation during each Congress to predict industry tariffs at the end of that period (e.g., majority representation over 1991–93 is used to predict tariffs in 1993).

A.3 Import Demand Elasticities

Import demand elasticities at the HS6 level are taken from Kee et al. (2008), as provided generously by Alessandro Nicita. These elasticities cover the period 1998–2001. We assign this elasticity to each associated HS10 industry (a concordance exists only from HS10 to SIC4). Using imports from 1993 (from Robert Feenstra's website) as weights, we then concord from HS10 to SIC4, using the concordance from Peter Schott's website. Any HS6 with missing elasticity is excluded from the industry total for calculating the weights, so that the weights sum to one for each industry. This yields import demand elasticity estimates for 374 manufacturing industries.

A.4 Organized Industries

For all classifications, we use contributions only to the majority party (Democrats in the first two electoral cycles in our sample, and Republicans in the third) for candidates running for the House of Representatives. Results do not differ substantially when we use contributions to all candidates; the correlation between contributions to Democrats and total contributions during the 1991-92 electoral cycle, for example, is 0.97. Note that contributions occur during the electoral cycles preceding the Representatives' taking office; for example, we use contributions during the 1989–90 electoral cycle to predict lobbying influence in the 102nd Congress (1991–92), which is tariff year 1993 in our sample.

G-M: For each electoral cycle, we identify significant breaks in the distribution of industry-level contributions and in the distribution of contributions scaled by industry shipments. We consider 4 possible classifications; the results used here are based on noticeable splits that resulted in approximately 60% of the sample being classified as organized in each period.

G-B: We use contributions to majority-party candidates scaled by industry value shipments in each electoral cycle. These (4-digit) industry-level contributions are regressed on 2-digit industry dummy variables interacted with bilateral import penetration data for each of five trading partners (France, Germany, Italy, Japan, and the U.K.). An industry is "organized" with respect to a trading partner if the coefficient estimate on the interaction term is positive. This is repeated for each of the trading partners; an industry is classified as "organized" if it is organized with respect to any trading partner.

E-M: All industries with positive contributions are considered organized. This procedure leads to almost all industries being classified as organized, as is evident from Table 4.

B Supplementary Results: State-Level Analysis

Constructing a majority variable based on states rather than districts as the regional unit allows us to include the Senate as well as the House of Representatives in our analysis (Senators, of course, are elected on a state-wide basis), and we think that it may also reflect some of the way party politics work in practice. For example, there is a substantial amount of give-and-take among elected officials from the same state, especially within a party. Particularly in more densely populated states, people frequently live (and vote) in one Congressional district and work in another; many representatives have aspirations to state-wide office (senator, governor) and so try to appeal to state residents outside of their immediate district; and many industries develop nearby suppliers, which may be located in a distinct district, but lead to correlated inter-district economic interests. For this reason, double-checking our results by constructing the majority variable by summing over majority states rather than majority districts appears a useful exercise.

There is a complication when we switch from districts to states, however: in the model, each district is represented by one legislator, and $(\alpha_{jg} - \beta_j)$ is summed over the majority districts. But states are represented by several legislators, who may belong to different parties, so it is not quite clear how to define what constitutes a majority state. In the specification presented below, we counted a state as a majority state if at least half of its legislators (calculated separately for the House and the Senate) came from the majority party. We then summed over majority states, and averaged the results for the House and the Senate. Comparing Tables 4 and 5, we see that the state majority variable exhibits less variation than the district majority variable; i.e., at the state level, industries are more evenly distributed than at the district level. However, there is not much difference between the means of the two specifications.

The results in Table 6 show that the qualitative results of the district-level speci-

Table 5: Additional Data Summary: State Specification

tariff equation 1993 (319 obs.)				
Variable	Mean	Std. Dev.	Min.	Max.
E-M lobby	.9937304	.0790562	0	1
G-M lobby	.5768025	.4948424	0	1
G-B lobby	.7774295	.4166258	0	1
$z_g = \sum_{j \in K} (\alpha_{jg} - \beta_j)$ state	-.0025	.0714	-.2804	.1568
$\sum_{j \in K} \beta_j$ state	.8396	–	–	–
tariff equation 1995 (314 obs.)				
Variable	Mean	Std. Dev.	Min.	Max.
E-M lobby	1	0	1	1
G-M lobby	.5955414	.4915703	0	1
G-B lobby	.5191083	.5004322	0	1
$z_g = \sum_{j \in K} (\alpha_{jg} - \beta_j)$ state	.0104	.0573	-.2261	.1458
$\sum_{j \in K} \beta_j$ state	.8542	–	–	–
tariff equation 1997 (316 obs.)				
Variable	Mean	Std. Dev.	Min.	Max.
E-M lobby	1	0	1	1
G-M lobby	.64874342	.478123	0	1
G-B lobby	.6360759	.4818902	0	1
$z_g = \sum_{j \in K} (\alpha_{jg} - \beta_j)$ state	.0174	.0743	-.2143	.2787
$\sum_{j \in K} \beta_j$ state	.6463	–	–	–

Table 6: Panel Estimation Results 1993 and 1995: State Specification

Lobby:	G-M	G-B	E-M
$\hat{\gamma}_1$	.0089***	.0088***	.0076***
	(.0011)	(.0014)	(.0012)
$\hat{\gamma}_2$	.0204*	.0296**	.0186*
	(.0114)	(.0132)	(.0107)
$\hat{\gamma}_{1993}$	66.50***	67.09***	69.43***
	(6.27)	(7.11)	(6.41)
$\hat{\gamma}_{1995}$	76.21***	89.71***	79.39***
	(7.79)	(9.67)	(7.92)
implied structural parameters			
$\hat{\phi}$	.0090***	.0091***	.0077***
	(.0011)	(.0015)	(.0012)
$\hat{\zeta}$	.0207*	.0304**	.0189*
	(.0118)	(.0139)	(.0110)
number of observations	633	651	633
first-stage F-stat lobby	13.43	18.91	47.95
first-stage F-stat majority	58.22	58.83	58.22
Hansen's J-stat p-value	.4274	.3145	.5502

Standard errors (in parentheses) for structural parameters were calculated using the delta method. *, **, and *** indicate significance at 10%, 5% and 1% levels. Instruments are industry employment, the majority variable z_g (calculated at the state level) and z_g interacted with industry capital.

Table 7: Predicted Tariff Changes Due to Changes in Lobby and State Majority Variable

specification:	G-M	G-B	E-M
lobby min to max	2.8833	2.8796	2.4455
majority min to max	2.8766	4.2134	2.6699
ratio maj to lobby	0.9977	1.4632	1.0918
lobby mean to mean + 1 std.	1.4483	1.3471	–
majority mean to mean +1 std.	0.4290	0.6319	0.3978
ratio maj to lobby	0.2962	0.4691	–

The predicted ad valorem tariff changes are given in percentage points and were calculated using the sample averages for 1993 and 1995 as reported in Table 4. The standard deviation for lobby and majority variable are calculated as weighted average over time rather than over the entire two-year sample in order to allow the reader to reconstruct the results using the provided data.

fication are confirmed. Both the majority and the lobby variable have positive signs and are statistically significant (the lobby variable at the 1% level, the majority variable at at least the 10% level), and the coefficient on the majority variable is higher than the lobby variable coefficient, although less so than in the district specification: The estimates of ζ lie between 1.89% and 3.04% compared to values between 6.18% and 7.44% in the district analysis.

When we look at the induced tariff changes when we change the lobby variable relative to the majority variable in Table 7, we see that because the majority variable coefficient is smaller in magnitude and in spread, the induced tariff changes are also smaller and in fact, end up being close to the lobby-induced changes when we go from the minimum to the maximum of the variable. They are considerably lower than the lobby-induced changes when we start from the mean and go up one standard deviation. Finally, just as in the district analysis, the statistical significance of the majority variable vanishes once the year 1997 is added (the full panel results were omitted here, but are available from the authors upon request).

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Notes

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[1]As quoted in "The Yale Book of Quotations", ed. Fred R. Shapiro, New Haven 2006.

[2]Since our analysis is at the industry level, this time period also avoids the change from the SIC to NAICS classification system in 1997.

[3]Gawande and Krishna (2003) provide an excellent survey of the empirical literature on the political economy of trade policy.

[4]See Persson and Tabellini (2003) for an extensive survey of the related public finance literature.

[5]Ansolabehere and Snyder (2006) argue that this spending strategy raises turnout.

[6]This is, we assume that minority delegates are not able to fully recompensate majority delegates as to induce maximization of general domestic welfare.

[7]G-B show that the level of protection is an increasing function of the tariff on an intermediate input. Since this issue has already been explored and because our focus is on the effects of a majoritarian electoral system, in this paper we abstract from issues involving intermediate goods and their industry lobbying.

[8]The other prediction of the G-H model is that $\gamma_0 = 0$. We are less concerned

with testing this, possibly over-restrictive, prediction.

[9]Since the RHS variables are lagged by approximately two years, the endogeneity issue may not be that important, however. Indeed, our instrumental variable results are quite close to the ordinary least squares results not reported in this paper.

[10]This exercise does not work well for the E-M lobby specification, because there is practically no variation in the lobby variable, i.e. almost all industries lobby.

[11]The majority variable coefficient is still positive, but small and never statistically significant. Full panel results are available from the authors upon request.

[12]An alternative interpretation would be that Republicans are less prone to using trade policy as a means to favor their own districts than Democrats. But to test such a hypothesis, we would need a much longer time series of data. Changing trade policy was clearly not a priority of the new Republican majority: The "Contract with America", on which many of them ran, consisted almost entirely of domestic policy items (the only exception involved the circumstances under which U.S. armed forces could serve under United Nations command). This does not necessarily mean that trade policy was not important in later years, only that it was not a top priority when the new Republican majority took office in 1995.

[13]More precisely, Democrats had been in the majority since 1987. Moreover, with a relatively short interruption from 1981 to 1987, the Democrats were the majority party since 1955, giving them ample time to shape trade policy.