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Antitrust and Trade Policy: Are Legislators Consistent?

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Abstract

In analyzing the potential exercise of monopoly power in domestic markets, both the academic literature and U.S. antitrust authorities have acknowledged the disciplining role of competition from abroad. In this paper, we explore the extent to which this view seems to reveal itself in recent U.S. Congressional votes taken during the 108th Congress (2003-04) on four free trade agreements (FTAs). To the extent that antitrust and trade liberalization are both viewed as pro-consumer in nature, we would expect to see a positive relationship between antitrust enforcement in their legislative district and Congressional votes in support of new FTAs. We do find evidence supporting a positive relationship between state-level antitrust enforcement (measured either by absolute number of cases filed or by cases relative to the state's economic size) and support for FTAs.

Key words: Antitrust, Trade Protection

JEL classification: F13, L4

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I. Introduction

It is increasingly the case that entry, either from domestic or foreign origins, is considered to be a major determinant of competition in economic markets. A considerable amount of empirical literature in Industrial Organization has found imports or international pressures measured in other ways (*e.g.*, exchange rate movements) to influence domestic prices and profits. In analyzing the potential exercise of monopoly power in domestic markets, both the academic literature and U.S. antitrust authorities have acknowledged the disciplining role of competition from abroad. This suggests that the discipline of trade liberalization may be viewed by legislators as consistent with traditional antitrust enforcement in promoting well-functioning domestic markets. In this paper, we explore the extent to which this view seems to reveal itself in recent U.S. Congressional votes taken during the 108th Congress (2003-04) on four free trade agreements (FTAs). To the extent that antitrust and trade liberalization are both viewed as pro-consumer in nature, we would expect to see a positive relationship between antitrust enforcement in their legislative district and Congressional votes in support of new FTAs.

II. Literature Review

One of the earliest studies addressing the impact of imports on the domestic industry was Esposito and Esposito (1971). Although their econometric results were quite weak and fragile with respect to model specification, more recent studies (*e.g.*, DeRosa and Goldstein (1981) and DeGhellinck *et al.* (1988)) have found similar disciplining impacts of foreign competition. Feinberg (1989a) suggests that imports –

especially from developing economies which have been the target of many recent FTAs – are likely to have especially strong impacts on domestic markets where tacit collusion may otherwise be possible. Feinberg and Shaanan (1997) find this to hold empirically and that an easing of non-tariff barriers (as expected to emerge through FTAs) would also likely have desirable pro-competitive impacts.[1]

The joint U.S. Department of Justice/Federal Trade Commission Merger Guidelines (as revised in 1997) say (in section 3.0):

> A merger is not likely to create or enhance market power or to facilitate its exercise, if entry into the market is so easy that market participants, after the merger, either collectively or unilaterally could not profitably maintain a price increase above premerger levels. ….In markets where entry is that easy (i.e., where entry passes these tests of timeliness, likelihood, and sufficiency), the merger raises no antitrust concern and ordinarily requires no further analysis.

While the guidelines do not specifically discuss supply from abroad it is understood that this competition is potentially relevant to understanding domestic market competition. It is also clear (from section 1.3 of the Guidelines) that foreign firms either currently selling in the U.S. market or which could quickly supply the U.S. market in response to price incentives are treated in the same manner as domestic firms in judging the nature of competition.

If it is the case that less-restrictive trade policy and more active domestic antitrust enforcement have similar impacts on domestic prices, we should expect legislators to vote in similar ways on the two types of policies. Unfortunately, votes on antitrust issues are quite rare. However, we argue that a good indication of how a legislator ***would*** vote on antitrust issues is the aggressiveness of antitrust enforcement in his/her home state.

[1] Similarly, Feinberg (1989b) provides evidence suggestive of exchange rate impacts on domestic prices being limited by non-tariff barriers to trade.

We incorporate a measure of this aggressiveness in probit models used to explain the recent votes by members of Congress on FTAs.[2]

The standard economic assumption underlying voting behavior is that representatives are concerned with electoral success (*e.g.*, reelection or election to higher office). As a consequence, the basic model is that representative votes are determined by the economic and ideological interests of the legislator's district. Baldwin (1985) identifies factors that would tend to increase support for "protectionist" policies – the greater the proportion of workers in the congressional district employed by import-sensitive industries and the smaller the proportion of workers in the district employed by export industries. Other determinants include the policy position of the member's political party, the preferences of the President (stronger if from the same political party), the member's congressional leadership, and general support for policies such as income assistance for low-income workers and retaliation against "unfair" trade practices.[3]

Empirical tests of these theories have been carried out on roll-call votes on several FTAs. The analysis is generally undertaken on the votes in one of the houses of Congress. Nollen and Iglarsh (1990) examine votes taken by U.S. Senators on an amendment to the *1984 Omnibus Trade Bill*, the *1985 Textile Import Quotas Bill*, and the *1987 Omnibus Trade Bill*. They incorporate measures of the state's measure of export dependence and indicators of the impact of imports on the state. To capture ideological influences, they also include measures of the legislator's political party as well as pro-

[2] This measure also has the advantage of being more clearly exogenous – in this sense it might be viewed as an instrument for the more endogenously determined vote on a (hypothetical) antitrust matter.

[3] Baldwin (1985) tested and found support for several of these theories focusing on U.S. House and Senate votes on the *Trade Act of 1974*.

business and pro-labor ratings. They find empirical support that legislator votes on trade bills are determined by constituent interest and ideology.

Marks (1993) analyzes congressional voting on five amendments to the omnibus trade bills of 1987. Explanatory variables in his models include the employment levels in six "trade-sensitive" industries – textiles, footwear, steel products, machine tools, semiconductors, and autos/automobile parts to account for the impact on affected constituent groups.

Kahane (1996) focuses on the votes in the U.S. Senate on the 1991 extension of fast-track procedures, a precursor to the 1993 congressional vote on the North American Free Trade Agreement (NAFTA). His model of Senate votes includes employment data for several industries affected by free trade with Mexico to measure the "winners" and "losers" of the fast-track authority, as well as measures of union membership, political party, and ideological measures. Overall, he found the larger were the groups identified as losers as a result of fast-track authority in the state, the less likely was the senator to vote for extending the authority.

Probably the most important FTA, at least as measured by public attention, was NAFTA. Several papers focus on the determinants of the roll-call votes on this bill.[4] Baldwin and Magee (2000) analyze congressional voting on NAFTA, along with votes on the General Agreement on Tariffs and Trade (GATT), and "most favored nation" (MFN) treatment for China. In addition to measures of union strength, income, and a measure of the legislator's ideology, they include employment levels from several industries determined to be "winners" and "losers" of these FTAs. The empirical results

[4] In addition to those discussed here, others published in the political science literature include Box-Steffensmeier *et al*. (1997), Holian *et al*. (1997), and Uslaner (1998).

from probit models on the votes for these trade measures indicate that employment measures in most of the individual industries, either "winners" or "losers," did not have much of a statistical impact. Kamdar and Gonzalez (1998) also study U.S. Senate votes on NAFTA and GATT. They include the output as a share of gross state product of certain industries identified as "winners" (primarily capital intensive industries and the service industry) and "losers" (the "low-skill" industries) as a result of these free trade measures. In addition, they include measures of unemployment change, labor union membership, the number of workers receiving trade adjustment assistance, the percentage of Hispanic voters in the state, measures of corporate and labor campaign contributions and membership on the Senate International Trade Subcommittee of the Finance Committee and the Foreign Relations Committee. The measures associated with the "loser" industries were the expected sign in the NAFTA vote while the "winner" industry measures were statistically insignificant in both the NAFTA and GATT.

Arce, Koopman, and Tsigas (2008) follow a similar empirical framework in their estimation of the determinants of the U.S. Senator's free-trade position, as measured by an index based on votes taken on ten trade bills during the 108th Congress. They find the Senator's political party, percentage of labor unionization in the state, and the amount of the Senator's political contributions from business affect the Senator's free-trade position. The authors do not use the standard industry measures in their regressions, instead employing a general equilibrium model to calculate the impact on Gross State Product (GSP) caused by changes in import/export prices. They find the presence of import industries such as food and textiles ("losers") to have a larger negative impact on the Senator's free trade position than did export industry measures.

A final note about the literature on the determinants of congressional voting on FTAs is whether there are differences between the outcome of roll-call votes taken in the Senate and the House of Representatives. As described, most of this literature focuses on only one house of Congress, so comparisons on the determinants of votes between the two houses of Congress is unclear. Baldwin (1985, p. 51) lists possible differences in characteristics of the two houses – "size, constitutional functions, and rules governing behavior" – as factors that could explain differences in the types of trade legislation that are voted on in each. These factors all suggest stronger support for FTAs in the Senate roll-call voting. Because it is a smaller institution, it is easier for members of the Senate to present their views and to negotiate with a potential majority of colleagues in the individual committees and on the Senate floor. Baldwin also notes that Senators may take more seriously the constitutional authority of Congress with respect to trade legislation; as a result, they might be more likely to pass trade legislation in order to assert some autonomy over the President's authority. Finally, the longer terms in office of Senators may shield them to some extent from special interest lobbying.

III. Econometric Specification and Data

Over the past decade there has been a great deal of attention given by policy makers to negotiating bilateral FTAs (in part as a substitute for additional multilateral trade liberalization which has floundered since the creation of the World Trade Organization in the mid-1990s). In this paper we focus on four such FTAs – with Australia, Chile, Morocco, and Singapore – which were voted on (and passed) by both houses of Congress during the 108th Congress (2003-04).

As reported in Table 1, each of the FTAs passed by a clear majority, although the agreements with Australia and Morocco enjoyed more support than the agreements with Singapore and Chile in both the House and the Senate. As suggested by the discussion above, all four agreements passed by a greater majority in the Senate when compared to the House. There is an extremely high degree of correlation (0.92) between legislator's votes on the Chilean and Singaporean FTAs, which were considered in the House and Senate on the same day, but less correlation among the votes on the other FTAs. For example, the degree of correlation between legislator's votes on the Australian and Moroccan FTAs, which were considered during the same month, is only 0.60.

Following the political economy literature described above, a legislator's decision whether to vote in favor of a FTA is a function of legislator and constituent characteristics, as well as other unobserved factors. It is likely that legislators who vote in favor of one FTA will also vote in favor of other FTAs, thus the residuals in the empirical analysis of votes on the Australian, Chilean, Moroccan and Singaporean FTAs will likely be correlated. To account for this fact, we estimate a multivariate probit model in which the legislator *i's* vote on FTA *j* ($Vote_{ij}$) will be Yea, or equal 1, according to the specification:

$$\begin{aligned} &Vote_{ij} = 1(X_i\beta_j + \varepsilon_{ij} > 0) \\ &E[\varepsilon_j] = 0 \\ &Var[\varepsilon_j] = 1 \\ &Cov[\varepsilon_j, \varepsilon_k] = \rho_{jk} \end{aligned} \quad . \qquad (1)$$

In this equation, X_i is a vector of constituent and legislator characteristics, and β_j are the parameters to be estimated. Note that the degree of correlation between the legislator's vote on FTA *j* and FTA *k* is captured by ρ_{jk}.

We expect that – controlling for other factors to be discussed below – legislators who represent districts which have a positive assessment of antitrust enforcement will be more likely to support these FTAs. Our measure of antitrust enforcement is the aggressiveness of state Attorneys General in pursuing (as lead plaintiffs) antitrust actions under both state and federal statutes; based on a database (maintained by the National Association of State Attorneys General) of cases filed between 1990 and 2006, we calculate both the absolute number of cases *(Cases)* filed as well this number relative to a state's economic size *(Cases / GSP)* as measured by its Gross State Product (GSP) in billions of dollars.[5]

We include a number of legislator-specific variables. Senators have a broader constituent base and longer terms; as a result, they may be less vulnerable than Representatives to narrow interest groups. To account for this possibility, we include a dummy variable for Senators *(Senate)* in some specifications and estimate the model separately for Senate and House members in other specifications. The Republican Party, in recent years, has been considered less protectionist than the Democratic Party. Therefore, we hypothesize that members of the Republican Party *(Republican)* are more likely to vote for the FTAs in this sample. We include the number of terms each legislator has served in office *(Terms)* to account for the possibility that more senior members of Congress are less likely to be influenced by interest groups associated with trade issues. Information on the members of Congress comes from Congressional Quarterly's Congress Collection Database.

[5] Results from specifications using the number of antitrust cases relative to the state's population were not significantly different from those presented here.

Other district economic characteristics are expected to affect the Congressional voting. Legislators from districts with extremely high unemployment rates may be less likely to vote for free trade agreements, in the fear that increased trade will exacerbate poor economic conditions. We include the average unemployment rate in 2003 in the Congressional district *(Unemployment Rate)* to account for this possibility. The Bureau of Labor Statistics (BLS) Local Area Unemployment Statistics provides state and county-level unemployment rates. We use county-level unemployment rates from the BLS to calculate a weighted average unemployment rate in the Congressional district using the concordance available from the Missouri Census Data Center's Geocorr2K website.[6]

Industries producing traded goods are more likely to be affected by trade liberalization than non-traded goods industries. Using employment data from the U.S. Census Bureau's 2003 County Business Patterns, we calculate each district's share of employment in the manufacturing sector (*Manufacturing)* as defined by their North American Industrial Classification (NAICS) code.[7]

Because labor unions typically oppose free trade agreements, we would expect the likelihood of voting for any of the agreements analyzed in this sample to decrease with the unionization rate of the state *(Unionization)*. We measure unionization using the percentage of state workers that are members of unions as reported in the Hirsch and Macpherson (2003) Union Membership and Coverage Database.

[6] Unemployment rates from each county were weighted by the percentage of the population in each Congressional district from that county prior to taking the Congressional district average.

[7] Similar variables measuring the share of employment in the agriculture and primary metal industries were statistically insignificant, thus excluded from the final specifications. We estimated the number of employees in Congressional Districts in the manufacturing (NAICS 31-33) sector by allocating county employment to various districts using the percentage of the county residing in each Congressional District as provided by Geocorr2K dataset discussed above. The share of employment in the manufacturing sector was then calculated by dividing by the estimated total employment in the district.

We exclude from the estimation legislators who neglected to participate in one or more of the votes on the FTAs considered in this sample, which leaves us with a dataset of the votes of 495 legislators (95 Senators and 400 Representatives) on all four FTAs considered during the 108th Congressional session. Summary statistics of the explanatory variables are included in Table 2.

Although political contributions from firms, labor interests and business interests often have proved to be statistically significant determinants of Congressional voting on trade legislation in the literature described in Section II, we omit political contribution data from our econometric model. As discussed by Baldwin and Magee (2000) and Liebman and Reynolds (2006), the same unobserved factors likely influence both political contributions to a legislator and his or her vote on passage of a new FTA; as a result, political contributions are endogenous to the model and inclusion of contributions variables would bias the results. There is unlikely to be much correlation between our primary variable of interest, state antitrust enforcement, and political contributions to the individual legislators from that state, thus omitting contributions should not bias the key results discussed below.[8]

IV. Results

Parameters from the estimation of the multivariate probit in which we control for the absolute number of antitrust cases filed in the Congressmen's state *(Cases)* are included in Table 3. The specification presented in column 1 constrains the impact of

[8] Moreover, because many of the same determinants affect both political contributions to the legislator and the legislator's vote on a particular piece of trade legislation, including the legislator's political party, unionization rate in the state, and the number of terms in office, we are implicitly instrumenting for political contributions.

Cases to be equal across all four of the FTAs in our sample, while the specification presented in column 2 allows the impact of *Cases* to vary across these FTAs.

As we hypothesized, the absolute number of antitrust cases filed in the Congressman's state has a positive and statistically significant impact on the likelihood of voting in favor of FTAs, suggesting that members from states with strong anti-trust enforcement are more likely to vote in favor of free trade agreements that can also increase competition and improve consumer welfare. The impact appears to be stronger on the likelihood of voting in favor of the FTAs with Australia and Singapore when compared to the agreements with Chile and Morocco; in fact the impact of *Cases* on the likelihood of voting in favor of the FTA with Morocco is statistically insignificant.

Table 4 uses the coefficient estimates to calculate the average marginal effect of a 10 percent increase in the explanatory variables on the likelihood of voting in favor of all four FTAs in the sample. A 10 percent increase in the number of antitrust cases filed in the state results in a 0.65 percentage point increase in the likelihood of voting in favor of all four FTAs. This represents a 1.2 percent increase in the average estimated likelihood of a favorable vote for all four FTAs of 53.05 percent. The empirical evidence suggests that legislators view both trade liberalization and antitrust enforcement as policies that promote well-functioning, competitive domestic markets. Thus, policy makers from states that value competition are more likely to vote in favor of new FTAs.

The final two columns of Table 3 presents the parameter estimates associated with the likelihood of voting for the four FTAs separately for House and Senate members. Because there are fewer observations, we are unable to estimate the four-equation multivariate probit model on the Senate sub-sample. Instead, we estimate four probit

models, one for each FTA. The empirical results suggest that the positive relationship between antitrust enforcement and votes in favor of trade liberalization is stronger among House members than among Senators. There is a positive and significant relationship of *Cases* on a House members vote in favor of FTAs with both Australia and Singapore; the relationship between *Cases* and House members votes on the Chilean and Moroccan FTAs is also positive, albeit statistically insignificant. As reported in Table 4, a 10 percent increase in the number of antitrust cases filed by a state results in an average increase of 0.58 percentage points in the likelihood of a House member voting in favor of all four FTAs; this represents an increase of slightly more than 1 percent in the estimated likelihood of 53.06 percent. There is weaker evidence of a positive relationship of *Cases* on the likelihood of a Senator voting in favor on an FTA; the coefficient on *Cases* is statistically significant only in the case of the Senate vote on the FTA with Chile.

Although there is a high correlation between the number of antitrust cases filed by the state and the economic size of the state, the statistically-significant positive impact of anti-trust enforcement on support for FTAs does not appear to be driven by this correlation. Results from unreported specifications that include GSP as a separate regressand were not qualitatively different from those presented here.

To further investigate whether the results from Table 3 are being driven by the correlation between antitrust enforcement and the economic size of the state, Table 5 presents the results from specifications that include the number of Cases filed per billion dollars of the state's GSP as well as the square of this term. Coefficients on both terms are significant and indicate that there is a positive and significant relationship between the relative degree of antitrust enforcement in the policy-maker's state and the likelihood

that the policy-maker votes in favor of new trade liberalization efforts. Specifically, we find that the likelihood of voting in favor of all four FTAs in our sample increases with the relative degree of antitrust enforcement, but at a decreasing rate.[9] Intuitively, even a modest amount of antitrust enforcement may suggest that the state and its elected policy-makers are pro-competition and thus would be in favor of new trade liberalization efforts, but the impact of further antitrust actions beyond some threshold is limited. Calculations similar to those presented in Table 4 reveal that a 10 percent increase in the intensity of state antitrust case filings (relative to economic size) would result in a 0.28 percentage point increase in the likelihood of voting in favor of all four FTAs, an increase in the estimated probability of approximately 0.5 percent. As illustrated in Columns 3 and 4 of the table, the positive relationship between antitrust enforcement and votes in favor of new trade liberalization is only statistically significant in the House.

Many of the parameter estimates associated with other explanatory variables are also of the hypothesized sign, and statistically significant in most of our specifications. As expected and found in earlier studies, Republicans are statistically more likely to vote in favor of increased trade liberalization, while those members from districts with high unionization rates are less likely to vote in favor of increased liberalization.

Theoretically, it is unclear what impact the share of manufacturing employment in the Congressional district should have on the likelihood of voting in favor of increased trade liberalization. Legislators from districts with more export-intensive manufacturing would be more likely to vote for new FTAs, while those from districts with more import-

[9] The parameter estimates from column 1 suggest that there is a positive relationship between state-level antitrust enforcement and the likelihood of voting in favor of FTAs as long as the number of antitrust cases filed per billion dollars of GSP is less than 0.54. This is true of all states but Maine. Coefficients on the square of the absolute number of Cases in specifications similar to those reported in Table 3 were insignificant.

sensitive manufacturing would be less likely to vote for new FTAs. Empirically, however, the share of manufacturing employment in the Congressional district proved to have a negative and significant impact on the likelihood of voting in favor of the FTAs in this sample, at least in most of our specifications.[10]

Surprisingly, the district unemployment rate is an insignificant determinant of Congressional votes on FTAs. This may indicate that Congressmen are not swayed by short-term economic conditions when making long-term policy decisions such as whether the United States should form a free trade agreement with another country.

Finally, with the possible exception of the Australia FTA vote, Senators do not appear to be more likely to vote in favor of new free trade agreements when compared to their counterparts in the House. Although the length of time that the Congressman has served in office is not a statistically strong determinant of the likelihood of voting for an FTA, there is evidence in some of the specifications that the likelihood of voting in favor of the FTAs with Chile and Singapore decreases with the number of terms the Congressmen have served in office.

V. Conclusions

The aggressiveness of enforcement of state and federal antitrust statutes by state Attorneys General is an indicator of the sentiment within that state in favor of consumer interests. As such, it would be expected that this sentiment would also be expressed by

[10] In other specifications not reported here we include the impact of a change in U.S. imports and U.S. exports on GSP as simulated in Arce, Koopman and Tsigas (2008). The *Import Sensitivity* variable represents the change in GSP that would occur if the rest-of-the-world's production costs fell by one percent, while the *Export Sensitivity* variable represents the change in GSP that would occur if the rest-of-the-world's willingness to pay for U.S. products increased by one percent. These variables proved to be insignificant.

the state's Congressional delegation in their votes on free trade agreements – which are also widely viewed as pro-competition and hence, pro-consumer. This study is the first to empirically examine this issue. We find evidence that increased state antitrust enforcement is associated with greater support for negotiated free trade agreements, especially for members of the U.S. House of Representatives.

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Table 1
Percentage Voting in Favor of U.S. FTA*

	House	Senate
Singapore *(July 24 and 31, 2003)*	63.7	67.3
Chile *(July 24 and 31, 2003)*	63.5	67.0
Australia *(July 14 and 15, 2004)*	74.2	83.3
Morocco *(July 22 and 21, 2004)*	76.5	86.7

*Dates in parentheses indicate the dates the House and Senate voted on the FTA, respectively.

Table 2
Summary Statistics

	Senate		House	
	Mean	Std. Dev.	Mean	Std. Dev.
Cases	9.032	13.722	19.630	19.110
Cases/GSP	0.056	0.104	0.046	0.058
Republican	0.526	0.502	0.525	0.500
Terms	3.347	1.773	5.755	3.997
Unemployment Rate	5.555	1.051	6.092	1.271
Manufacturing	0.126	0.046	0.131	0.633
Unionization	0.116	0.056	0.130	0.062
Number of Observations	95		400	

Table 3
Determinants of FTA Voting Behavior (Coefficients)--Cases*

	Constrained	*Unconstrained*	*House*	*Senate*
Australia				
Cases	**0.011 (3.58)**	**0.018 (3.95)**	**0.021 (4.13)**	-0.013 (0.84)
Senate	**0.342 (1.84)**	**0.382 (2.05)**		
Terms	-0.023 (1.28)	-0.023 (1.27)	-0.013 (0.72)	-0.101 (0.95)
Republican	**1.163 (7.64)**	**1.185 (7.71)**	**1.171 (6.82)**	**0.976 (2.20)**
Unemp. Rate	0.055 (0.96)	0.039 (0.66)	-0.023 (0.36)	**0.835 (3.04)**
Unionization	**-3.003 (2.47)**	**-3.699 (2.88)**	**-3.620 (2.58)**	**-10.155 (2.09)**
Manufacturing	**-2.233 (1.84)**	-1.724 (1.39)	-1.392 (1.59)	-6.420 (1.28)
Chile				
Cases	**0.011 (3.58)**	**0.009 (2.14)**	0.005 (1.19)	**0.028 (2.03)**
Senate	0.162 (0.99)	0.145 (0.87)		
Terms	**-0.030 (1.65)**	**-0.030 (1.65)**	-0.012 (0.62)	**-0.195 (2.11)**
Republican	**1.464 (10.4)**	**1.455 (10.3)**	**1.588 (9.73)**	**1.010 (3.03)**
Unemp. Rate	-0.017 (0.32)	-0.012 (0.22)	0.013 (0.22)	0.013 (0.08)
Unionization	**-3.218 (2.84)**	**-2.996 (2.60)**	**-2.604 (2.02)**	**-6.750 (1.97)**
Manufacturing	**-2.103 (1.87)**	**-2.324 (2.04)**	**-2.869 (2.32)**	1.412 (0.40)
Morocco				
Cases	**0.011 (3.58)**	0.006 (1.46)	0.007 (1.48)	0.034 (1.21)
Senate	0.279 (1.57)	0.247 (1.39)		
Terms	-0.023 (1.28)	-0.023 (1.26)	-0.003 (0.17)	-0.114 (1.18)
Republican	**1.095 (7.35)**	**1.080 (7.30)**	**1.130 (6.62)**	0.415 (1.07)
Unemp. Rate	-0.084 (1.49)	-0.067 (1.18)	-0.076 (1.21)	0.050 (0.26)
Unionization	-0.549 (0.46)	-0.232 (0.19)	-0.934 (0.70)	-0.598 (0.16)
Manufacturing	**-3.647 (3.18)**	**-4.069 (3.48)**	**-3.342 (2.61)**	**-7.045 (1.71)**
Singapore				
Cases	**0.011 (3.58)**	**0.010 (2.41)**	**0.010 (2.23)**	0.011 (0.88)
Senate	0.222 (1.45)	0.213 (1.38)		
Terms	-0.023 (1.28)	-0.024 (1.36)	-0.008 (0.42)	-0.134 (1.55)
Republican	**1.095 (7.35)**	**1.460 (10.7)**	**1.527 (9.54)**	**1.075 (3.35)**
Unemp. Rate	-0.084 (1.49)	0.018 (0.35)	0.035 (0.59)	0.150 (0.94)
Unionization	-0.549 (0.46)	**-2.523 (2.25)**	**-3.049 (2.41)**	-2.589 (0.83)
Manufacturing	**-3.647 (3.18)**	**-2.141 (1.94)**	**-2.145 (1.81)**	-3.313 (0.91)
No. of Obs.	495	495	400	95

* Absolute value of z-statistics in parentheses. Estimates of constant term not reported. Values in bold are significant at least at the 10 percent level.

Table 4
Marginal Effects of FTA Voting Behavior*

	Constrained	*Unconstrained*	*House*
Cases	0.6522	0.6095	0.5765
Senate**	7.6736	7.5260	
Terms	-0.4379	-0.4350	-0.1798
Republican**	51.1119	51.0659	52.9254
Unemp. Rate	0.0228	0.0443	-0.0602
Unionization	-1.1530	-1.1639	-1.1799
Manufacturing	-0.9952	-1.0187	-1.0481
$E(Pr(Vote_i=1, \text{all } i))$	53.0563		

* Average marginal effect of a ten percent increase in the explanatory variable on the likelihood of voting in favor of all four FTA agreements.

** Average marginal effect of a discrete change in the dummy variable from 0 to 1 on the likelihood of voting in favor of all four FTA agreements.

Table 5
Determinants of FTA Voting Behavior (Coefficients)—Cases per GSP*

	Constrained	*Unconstrained*	*House*	*Senate*
Australia				
Cases/GSP	**4.066 (2.71)**	**3.744 (1.74)**	**5.506 (2.12)**	-6.680 (1.35)
$(Cases/GSP)^2$	**-7.495 (2.78)**	**-8.201 (2.20)**	**-9.331 (1.99)**	5.655 (0.71)
Senate	0.185 (1.02)	0.200 (1.09)		
Terms	-0.025 (1.40)	-0.024 (1.37)	-0.017 (0.90)	0.000 (0.00)
Republican	**1.156 (7.61)**	**1.152 (7.53)**	**1.167 (6.84)**	**1.189 (2.23)**
Unemp. Rate	0.070 (1.25)	0.067 (1.19)	0.023 (0.39)	**0.876 (3.02)**
Unionization	**-2.462 (2.09)**	**-2.405 (2.02)**	-2.097 (1.63)	**-11.149 (2.17)**
Manufacturing	**-2.727 (2.29)**	**-2.795 (2.33)**	**-2.583 (2.05)**	-5.085 (1.02)
Chile				
Cases/GSP	**4.066 (2.71)**	**3.140 (1.66)**	1.201 (0.44)	-1.897 (0.42)
$(Cases/GSP)^2$	**-7.495 (2.78)**	-5.599 (1.59)	-7.465 (0.74)	7.462 (0.55)
Senate	-0.037 (0.24)	-0.040 (0.26)		
Terms	**-0.034 (1.90)**	**-0.034 (1.87)**	-0.016 (0.87)	**-0.211 (2.31)**
Republican	**1.435 (10.3)**	**1.424 (10.2)**	**1.584 (9.56)**	**0.811 (2.49)**
Unemp. Rate	-0.012 (0.22)	-0.008 (0.16)	0.022 (0.38)	0.103 (0.66)
Unionization	**-2.400 (2.19)**	**-2.284 (2.06)**	-1.874 (1.50)	-5.039 (1.61)
Manufacturing	**-2.298 (2.09)**	**-2.369 (2.15)**	**-3.417 (2.84)**	-0.381 (0.11)
Morocco				
Cases/GSP	**4.066 (2.71)**	**4.222 (2.01)**	**4.364 (1.74)**	-0.107 (0.02)
$(Cases/GSP)^2$	**-7.495 (2.78)**	**-7.060 (1.86)**	-7.509 (1.53)	1.962 (0.20)
Senate	0.107 (0.60)	0.103 (0.58)		
Terms	-0.027 (1.50)	-0.027 (1.47)	-0.010 (0.55)	-0.130 (1.33)
Republican	**1.067 (7.02)**	**1.070 (6.96)**	**1.217 (7.07)**	0.368 (0.94)
Unemp. Rate	-0.072 (1.28)	-0.069 (1.22)	-0.063 (1.03)	0.142 (0.76)
Unionization	0.031 (0.03)	-0.018 (0.02)	-0.241 (0.19)	0.799 (0.23)
Manufacturing	**-3.967 (3.45)**	**-3.942 (3.41)**	**-3.606 (2.95)**	**-8.091 (1.98)**
Singapore				
Cases/GSP	**4.066 (2.71)**	**4.698 (2.60)**	**6.417 (1.65)**	-4.591 (0.54)
$(Cases/GSP)^2$	**-7.495 (2.78)**	**-7.436 (2.17)**	**-45.843 (1.82)**	31.824 (0.81)
Senate	-0.033 (0.22)	-0.045 (0.30)		
Terms	**-0.034 (1.94)**	**-0.034 (1.91)**	-0.015 (0.81)	**-0.173 (1.91)**
Republican	**1.435 (10.4)**	**1.442 (10.4)**	**1.568 (9.93)**	**1.052 (3.18)**
Unemp. Rate	0.027 (0.50)	0.029 (0.54)	0.036 (0.61)	0.217 (1.35)
Unionization	**-1.803 (1.65)**	**-1.986 (1.79)**	**-2.409 (1.91)**	-1.573 (0.50)
Manufacturing	**-2.481 (2.28)**	**-2.427 (2.23)**	**-2.927 (2.49)**	-4.852 (1.32)
No. of Obs.	495	495	400	95

* Absolute value of z-statistics in parentheses. Estimates of constant term not reported. Values in bold are significant at least at the 10 percent level.