Votes, Vetoes, and Preferential Trading Agreements

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ABSTRACT

Preferential trading agreements (PTAs) are proliferating rapidly. By 2006, according to the World Trade Organization (WTO), nearly 300 PTAs were in force, covering approximately half of the overseas trade conducted worldwide. Although just about every country now belongs to a PTA, some states have rushed to join many of these arrangements, whereas others have joined very few of them. What explains these variations? Some studies have emphasized that states enter PTAs to generate economic gains. There is considerable evidence that PTAs have ambiguous welfare implications, which sheds doubt on the claim that countries join them for economic reasons alone. Instead, we emphasize the domestic political benefits and costs for leaders contemplating membership. Two domestic political factors—the nature of the regime and the number of veto players—play a significant role in determining whether countries sign a PTA. The results of our statistical tests furnish considerable support for these arguments. Based on an analysis of all PTAs formed since World War II, more democratic states are more likely to establish PTAs than their less democratic counterparts. We also find that states are less likely to enter a trade agreement as the number of veto players increases.

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from the demands of special interests. Leaders can help address this problem by entering a PTA. Since this problem is more severe in more competitive electoral systems, democratic chief executives are especially likely to join preferential arrangements. Second, we argue that leaders face transaction costs when making a trade agreement. The domestic ratification process contributes heavily to the magnitude of these costs. As the number of "veto players" expands, domestic ratification of an international agreement becomes more difficult. These two different domestic political factors—the nature of the regime and the number of veto players—play a significant role in determining whether countries are willing and able to establish a PTA.

The results of our statistical tests furnish considerable support for these arguments. Based on an analysis of all PTAs formed since World War II, we find that more democratic states are more likely to establish PTAs than their less democratic counterparts. We also find that states are less likely to enter a trade agreement as the number of veto players increases.

What are Preferential Trading Arrangements?

PTAs are international agreements that aim to promote economic integration among member-states by improving and stabilizing the access that each member has to the other participants' markets. There are five different types of PTAs (see generally: Bhagwati and Panagariya 1996; de Melo and Panagariya 1993; Pomfret 1988). First, some arrangements grant each participant preferential access to select segments of the other members' markets. Second, a free trade area (FTA) is marked by the elimination of trade barriers on many (if not all) products

In most countries, the executive branch sets the agenda in foreign affairs and has the power to initiate foreign economic policy. However, veto players must ratify policy choices made by the executive, such as joining a trade agreement. Formally, the head of state in a democracy—the prime minister, president, chancellor, or premier—is often required by the national constitution to obtain the approval of the legislature for international agreements, including PTAs. He or she will therefore need to anticipate the legislature's (or any other veto point's) reaction to the proposed arrangement and ensure it is domestically acceptable.

Ratification can also be less formal. In dictatorships, shifts in foreign economic policy

agreement that its terms would be weakened where domestic groups opposed it, or the government could exclude all sensitive sectors which affected veto players. In this way a government could negotiate any agreement so that its veto players did not oppose it. Research has suggested that governments do indeed respond to domestic conditions when designing international agreements (Downs et al. 1996; Koremenos et al. 2001; Rosendorff and Milner 2001). However, there are several constraints on such behavior. First, the government cannot expect to successfully negotiate whatever terms its domestic veto players want, since foreign countries have to sign on to the agreement. They are likely to want exactly those concessions that the domestic veto players oppose most fiercely. Second, as the number of veto players increases, the demands of these groups for exclusions or flexibility must also grow, making it more difficult for the executive to find an acceptable agreement with its foreign partners. Hence, as the number of veto players rises, the transaction costs of concluding an international agreement are likely to grow, and the possibility of forming a Pr negr nen eoti , -2.3 Td (200K(eot)-2(i)-11(, -

whether to enter into international negotiations and the terms of any prospective PTA prior to sitting down with foreign governments. By lobbying the government, interest groups can influence the government's bargaining position in negotiations over a PTA. The executive's position ex ante already reflects the influence of politically important interest groups.

Besides influencing the preferences of heads of state, interest groups have indirect effects through veto players. The distributional consequences of PTAs generate pressure for interest groups to organize and lobby for or against membership. In the trade policy literature, there is a long tradition of associating parties with the trade policy preferences of different interest groups (Rogowski 1989; Milner and Judkins 2004)

We begin by estimating the following model:

(1)

one observation corresponding to state i and a second observation corresponding to state j. For example, in the case of the United States-Canada dyad in 1985, we include one observation where the US is i and Canada is j, and a second observation where Canada is i

11 to each value, resulting in a range from 1 (highly autocratic) to 21 (highly democratic).

There are three principal reasons to rely on this measure in our empirical analysis. First, our argument treats regime type as a continuous variable, with the competitiveness of elections ranging from perfectly competitive to completely uncompetitive. As noted above, the index developed by Jaggers and Gurr has a range of 21 points, unlike some other measures that treat regime type as a categorical variable (Przeworski et al. 2000). Second, Jaggers and Gurr's index highlights a number of institutional dimensions of regime type that we stress. The ability of voters to choose the chief executive, which is central to our theory, is expected to rise as the process for selecting the executive becomes more competitive; as that process becomes more open; and as political participation becomes increasingly competitive. Jaggers and Gurr's index captures each of these three institutional elements, whereas various alternative measures do not (Gastil 1980 and 1990). Third, their index covers more countries during the period since World War II than most other measures of regime type (Gastil 1980 and 1990; Bollen 1980; Gasiorowski 1996).

From the standpoint of testing our theory, the second major independent variable in equation (1) is VETO PLAYERS_i. This variable, which is measured in year t, indicates the extent of constitutionally mandated institutions that can exercise veto power over decisions in state i as well as the alignment of actors' preferences between those institutions within the state. The data

Henisz's measure is well-suited to testing our theoretical model. The index is theoretically derived from a spatial model of veto players. The theory underlying his measure is very similar to our theory: it is a s

the preferences within those arrangements, it is appropriate for testing our hypotheses.

Control Variables

We also include a number of variables that previous studies have linked to the formation of PTAs to ensure that any observed effects of regime type or veto players are not due to other international or domestic factors. All of these variables except former colony, distance, and Post Cold War are lagged one year. Some of these variables also help us to control for differences in preferences between countries. For instance, countries without ongoing disputes and ones that are allies or that trade extensively may be much more likely to make agreements since they tend to share political and economic interests. Holding these factors constant is important for testing our argument about the effect of domestic political factors. We also need to account for a set of systemic factors that previous studies have linked to PTA formation.

First, TRADE_{ij} is the logarithm of the total value of trade (in constant 2000 US dollars) between countries i and j in year $t-1$.[3]

therefore use the PRIO data on interstate armed conflict, which covers the period from 1950 to 2005.[5] Just as disputes may inhibit PTA formation, close political-military relations may promote it (Gowa 1994; Mansfield 1993). We therefore analyze ALLY_{ij}, which equals 1 if countries i and j are members of a military alliance in year t-1, 0 otherwise. We code this variable using the Alliance Treaty Obligations and Provisions (ATOP) data (Leeds et al. 2002).[6] To ensure that our results are robust to the measures of disputes and alliances that are used, however, we conduct some additional tests after using the MIDs data to measure disputes and the Correlates of War (COW) data to measure alliances. Since previous research has found that a former colonial relationship between i and j increases the likelihood that they will enter the same PTA, we also include $\text{FORMER COLONY}_{ij}$, which equals 1 if countries i and j had a colonial relationship that ended after World War II, 0 otherwise (Mansfield et al. 2002, 499-501; Mansfield and Reinhardt 2003, 849-852).[7] Adding these variables allows us to account for some international factors that affect interstate commercial relations while analyzing the domestic

[5] We use v4-2008 of the data, available at http://www.prio.no/CSCW/Datasets/Armed-Conflict/UCDP-PRIO/old-versions/4-2007/. Their data includes four types of conflict: (1) extra-systemic armed conflict occurs between a state and a non-state group outside its own territory; (2) interstate armed conflict occurs between two or more states; (3) internal armed conflict occurs between the government of a state and one or more internal opposition group(s) without intervention from other states; and (4) internationalized internal armed conflict occurs between the government of a state and one or more internal opposition group(s) with intervention from other states (secondary parties) on one or both sides. The third type of conflicts was dropped. We retained the other three types and expanded the data so that all possible dyads between the countries on side A and those on side B of each conflict were included (Gleditsch et al. 2002, 7).

[6] For the ATOP data, we use version 3.0, specifically the atop3_0ddyr.dta file. Because the data end in 2003, we use data from 2003 to fill in 2004.

[7] Data on former colonial relations are taken from Kurian (1992).

PTA ratification.

Various studies have concluded that PTA formation is marked by a diffusion process, whereby the decision by one set of countries to join a preferential arrangement may lead others to do likewise (de Melo and Panagariya 1993, 5-

Some have claimed that countries are more likely to sign these before a round begins rather than during to increase their bargaining leverage (Crawford and Fiorentino 2005, 16; OECD 2001, 7). Others suggest that it is during the round itself that signing a PTA is most useful in this regard (Mansfield and Reinhardt 2003). We also explore whether the length of time since the last GATT/WTO round concluded might affect PTA formation. Some have argued that a long time between rounds might induce countries to seek PTAs (e.g., Fiorentino et al. 2007; Katada and Solis 2008).

We also include regional fixed effects, using the eight regional categories identified by the World Bank.[10] Finally, η_{ij} is a stochastic error term.

Descriptive statistics for all of these variables are presented in Table 1. The sample in the following analyses is comprised of all pairs of states during the period from 1950 to 2005. Because the observed value of the dependent variable is dichotomous, we use logistic regression to estimate the model. Tests of statistical significance are based on robust standard errors clustered on the dyad to address any heteroskedasticity in the data, as well as other problems associated with the directed dyad research design. To account for temporal dependence in the formation of PTAs, we include a spline function of the number of years that have elapsed (as of t) since each dyad last formed a PTA, with knots at years 1, 4 and 7, as suggested by Beck, Katz and Tucker (1998). In the following tables, however, the estimates of this function are omitted to conserve space.

[10] The World Bank does not give a region for advanced industrial countries, such as those in Western Europe, as well as the United States, Canada, and Japan. We assign them to their appropriate geographic regions. We also deviate from the World Bank in distinguishing South America from Central America and the Caribbean.

Results of the Empirical Analysis

fixed effects, indicating that within dyad variation in regime type and veto players has an especially potent impact on PTA ratification. The size of these effects is smallest when these variables are measured in year t-1, rather than year t. As we mentioned earlier, this is to be expected since our argument is that both factors should have an immediate effect on PTA ratification. Furthermore, whether we measure alliances and disputes using the ATOP and PRIO data or the COW data has little bearing on the estimated effects of regime type and veto players. Figures 1 and 2 show the effects of regime type and veto players, respectively, on the probability of PTA ratification based on the results in the first column of Table 2 (holding constant the remaining variables in the model).

To further illustrate the magnitude of these effects, we initially calculated the "relative risk" of state i ratifying a PTA with state j if the former state is democratic or if it is autocratic. More specifically, this risk is the predicted probability of state i e()3(hTw [(e()3(hTw [(e8oe)-20(iT.(r)5(-zTj

[Figures 1 and 2 about here]

Next, we compare the predicted probability of state i forming a PTA when it has few veto players –

PTAs are especially likely to form when hegemony erodes. The estimated coefficients of Hegemony are negative and statistically significant, indicating that the odds of ratifying a preferential arrangement rise as the portion of the world's output accounted for by the leading economy declines, except when dyad-specific fixed effects are introduced in the model. This

The results also show that contiguous states are unlikely to form PTAs. In combination with the observed effects of Distance_{ij}, this suggests that PTAs are most likely to form between states that are nearby but do not share a border. Finally, while many observers assume that PTAs are formed between a large, rich country and a small, poor one, our results indicate otherwise. The coefficient estimate of GDP Ratio_{ij} is negative and statistically significant, implying that greater imbalances in national income discourage the ratification of PTAs. Since countries that are equally powerful may be better able to conclude agreements that involve reciprocal concessions, this result may not be that surprising. But the idea that most small countries are forced into PTAs with larger ones against their will does not seem to be borne out (Gruber 2000).

The effects of alliances, GATT/WTO membership, existing PTA membership, contiguity, hegemony, and the GDP ratio change in key ways when we introduce fixed effects in the model. This undoubtedly reflects the fact that the vast bulk of dyads (fully 84 percent of the country-pairs in our sample) never form a PTA. These dyads are not used to generate the parameter estimates because introducing fixed effects leads us to assume that none of the independent variables in our model except for the fixed effects influences the probability of these dyads ratifying a PTA. It is for this reason that Beck and Katz (2001, 487-99; King et al. 2001) warn that using fixed-effects models to analyze time-series cross-section data with a binary dependent variable is "pernicious" and yields "estimates that are so far off as to be completely useless." This problem is exacerbated in our case because some independent variables – such as alliances and contiguity – display little change over time, even among the dyads that do form PTAs. These dyads are also excluded when estimating the coefficients of such variables. The upshot is that, given the sparseness of our data, very few dyads are used to estimate the model's parameters when including fixed effects, and fewer still are used to estimate some coefficients.

There is no reason to expect that these few pairs are a representative sample of the population of all dyads. Consequently, while we have included some results based on a fixed-effects specification because certain studies advocate this modeling strategy (Green et al. 2001), we think it is prudent to view the results in columns three and four only with the greatest caution and to place primary emphasis on the remaining results in Table 2.

Our analysis shows that a wide variety of factors influence PTA formation. Yet, even after accounting for domestic economic conditions, regional factors, and international influences, we find strong evidence that regime type and veto players shape the political calculus of governments contemplating PTAs. Our argument is not that the effect of domestic politics is

Henisz's (2000b, 2002) data and his measure of these players. As we discussed earlier, Beck et al. (2001, 2005) have developed an alternative measure, Checks, which emphasizes the extent of electoral competition, the number of domestic institutions that can check the chief executive, and the partisan differences across these institutions. To assess the robustness of our findings with respect to the measure of veto players, we replace Veto Players with Checks. As a result, this analysis spans the period 1975-2004 since Beck et al. do not provide data prior to 1975. In the first row of Table 3, we present the estimates of Checks and Regime Type based on this analysis. The remaining variables from the baseline model were included in this analysis, but are not presented in the table to conserve space and because there are few differences between their coefficient estimates in Table 2 and based on this analysis.

[Table 3 about here]

The results continue to indicate that states are increasingly unlikely to ratify PTAs as the number of veto players rises, since the estimated coefficient of Checks is negative and statistically significant. However, the quantitative influence of this variable is even larger than that of Veto Players. We again compare the predicted probability of state i forming a PTA when it has few veto players – which we define as the 10^{th} percentile in the data – to the predicted probability when it has many such players – which we define as the 90^{th} percentile in the data, holding constant the remaining variables in the model. Based on this analysis, a state with few Checks is about 55 percent more likely to ratify a PTA than one with more Checks.

Second, it is useful to address whether our results are sensitive to the particular measure of regime type that we are using. While Regime Type is particularly well-suited to testing our argument, the Polity Project has also developed an indicator of whether a given state holds competitive elections (Marshall and Jaggers 2005, 22). It is this feature of democracy that is

central to our argument because such elections increase the risk that leaders will be removed from office because of economic conditions that are actually beyond their control, thereby contributing to their interest in entering a PTA. We therefore replace Regime Type$_{it}$ with Competitive Elections$_{it}$, which equals 1 if state i is coded by the Polity Project as holding competitive elections as of year t, 0 otherwise.[12] As shown in the second row of Table 3, the estimated coefficient of this variable is positive and statistically significant. It is also large. Holding constant the remaining variables in the model, states with competitive elections are approximately 27 percent more likely to join a PTA than other countries. Furthermore, as shown in the third row of this table, the estimated coefficient of Competitive Elections$_{it}$ remains positive and significant (albeit smaller) even if we also include Regime Type$_{it}$ and both dyad-specific and year-specific fixed effects. These results provide a very strong support for our argument that it is competitive elections that underlie the decision by democracies to (o)2(s)1(410(y)22(5-6.8(mp)2(e)6(to)2((Tc(

To further address the effects of regime type, we undertake a number of analyses. We begin by recoding Regime Type$_{it}$ after excluding the constraints on the chief executive – which is one of the institutional features used to measure it – because these constraints may be closely related to the veto players that exist in a country. We would like to avoid including factors related to veto players in our measure of regime type. Next, we recode Regime Type$_{it}$ as a dichotomous variable. Consistent with various studies, we consider state i to be democratic in year t and assign it a score of 1 if Regime Type$_{it}$ • 2 W K H U Z L V H Z H F R Q V L G H U W democratic and assign it a value of 0. We also analyze a different dichotomous measure of regime type developed by Adam Przeworski and his colleagues (Przeworski et al. 2000). The

results of these analyses are presented in the fourth, fifth, and sixth row of Table 3. They continue to provide strong evidence that democracy promotes the ratification of PTAs since the coefficient estimates of Regime Type_{t-1} remain positive and statistically significant.

Third, we analyze whether our results are being driven by the European Community (EC) and the European Union (EU), institutions that are composed of democratic members. We find,

distinguish between situations in which no trade was conducted by a pair of countries and cases where the pair did not report any trade to the IMF. As such, it is not clear how to interpret situations in which the flow of trade is zero in the data set. Moreover, those dyads that actually did not conduct any trade in a given year could be considered unimportant to the international trading system and particularly unlikely to form any type of PTA. In the tenth row of Table 3, we report the estimated coefficients of

coefficients of these variables in table 5, as well as the coefficients of Regime Type$_{it}$ and Veto Players$_{it}$; but we do not report the coefficients of the remaining variables to conserve space. The results provide further evidence that PTA formation is guided by diffusion. As a country's regional neighbors form PTAs (Regional PTA$_{it}$), the likelihood rises that it will enter one too, suggesting that diffusion pressures exist at the regional level. Furthermore, as a country's major trading partners sign more PTAs (Trade Partner PTA$_{it}$), the country grows more likely to ratify such an arrangement.

[Table 5 about here]

To further analyze this issue, we follow Baldwin and Jaimovich (2010) by including Contagion$_{ij}$ and Contagion$_{ij}^2$ in our model. They find evidence of an inverted U-shaped relationship between their measure of contagion and PTA formation. As shown in the fourth column of Table 5, we find that this relationship is U-shaped since the estimated coefficient of Contagion$_{ij}$ is negative and the estimated coefficient of Contagion$_{ij}^2$ is positive. The differences between Baldwin and Jaimovich's results and ours may stem from the fact that their study covers fewer (113) countries and a shorter time frame (1977-2005). It may also stem from collinearity between Contagion$_{ij}$ a. .29.08 Tm (ij)Tj 0 Tc 0 1(A)2 Tm (ij)Tl(.)2 (23yt(s)-2(s)c11ht,t5hD0 8.04(a)Tj (go 0)2 c 0.a.o57Tl(.t) Tw [T0020. .29.08 Tm (ij

raises the specter of a simultaneity bias that complicates any effort to draw firm conclusions about the effects of diffusion and contagion pressures on the formation of preferential groupings (Baldwin and Jaimovich 2010, 12). Of central importance for present purposes, however, is that adding these three measures of diffusion has no bearing on the observed effects of regime type or veto players. Regardless of which measure is analyzed, the estimated coefficient of Regime Type$_{it}$ is positive, that of Veto Players$_{it}$ is negative, and both of them are statistically significant.

Ninth, we include a variable indicating whether state i was a post-Communist regime, as of year t-1. Various countries in Eastern Europe and the former Soviet Union rushed to enter PTAs during the 1990s and 2000s and we want to ensure that these states are not driving our results. However, this is not the case. The estimated coefficient of Communist$_{it}$ is not statistically significant and including this variable has very little bearing on our other results (Kornai 1992; US Central Intelligence Agency ; US Department of State).

Tenth, we examine whether accounting for the similarity of foreign policy preferences between states i and j affects our results. Signorino and Ritter (1999) argue that the S score$_{ij}$, which is a measure of the similarity of UN voting patterns between states i and j, provides a reliable estimate of the extent of such similarity. As such, it is not surprising that the coefficient of this variable is positive and statistically significant, indicating that states with more similar foreign policy preferences are especially likely to enter into PTAs. But adding this variable has no impact on the observed effects of regime type and veto players.

Thus far, we have treated the effects of regime type and veto players as linear. Finally, we analyze whether these factors have an interactive effect on PTA ratification. In fact, we find no evidence of this sort. When we add Regime Type$_{it}$ × Veto Players$_{it}$ to the baseline model, its estimated coefficient is neither large nor statistically significant. The results indicate that

regardless of the number of veto players, more democratic countries have a greater probability of ratification than less democratic states. Regardless of the country's regime type, a rising number of veto players always reduces the

economic ones, for leaders.

Second, one of the domestic impediments to entering a PTA is the transaction costs associated with ratifying the agreement. Trade accords involve the exchange of market access among countries. Some agreements also aim to coordinate members' trade regimes. These policy changes have domestic consequences. Certain groups gain from these barrier reductions; other groups lose. If these distributional losers have political clout, they can delay or block such policy change.

Veto players represent political interests other than the leader's party and have the institutional capacity to prevent change. Assuaging these groups can be time consuming and expensive. Leaders may have to alter the trade policy changes they would prefer and they may have to bribe veto players to gain their acquiescence. The more veto players that exist, therefore,

Table 1: Summary Statistics

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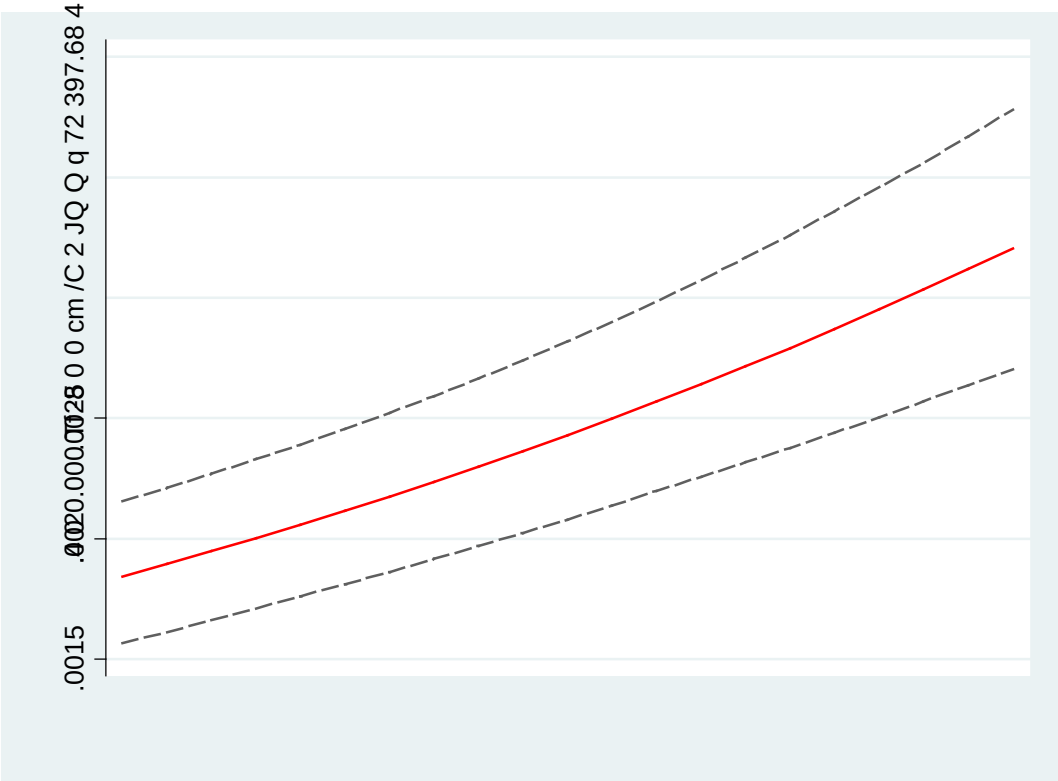
Table 3: Supplemental Tests of the Effects of Regime Type and Veto Players on PTA Ratification, 1951-2004.

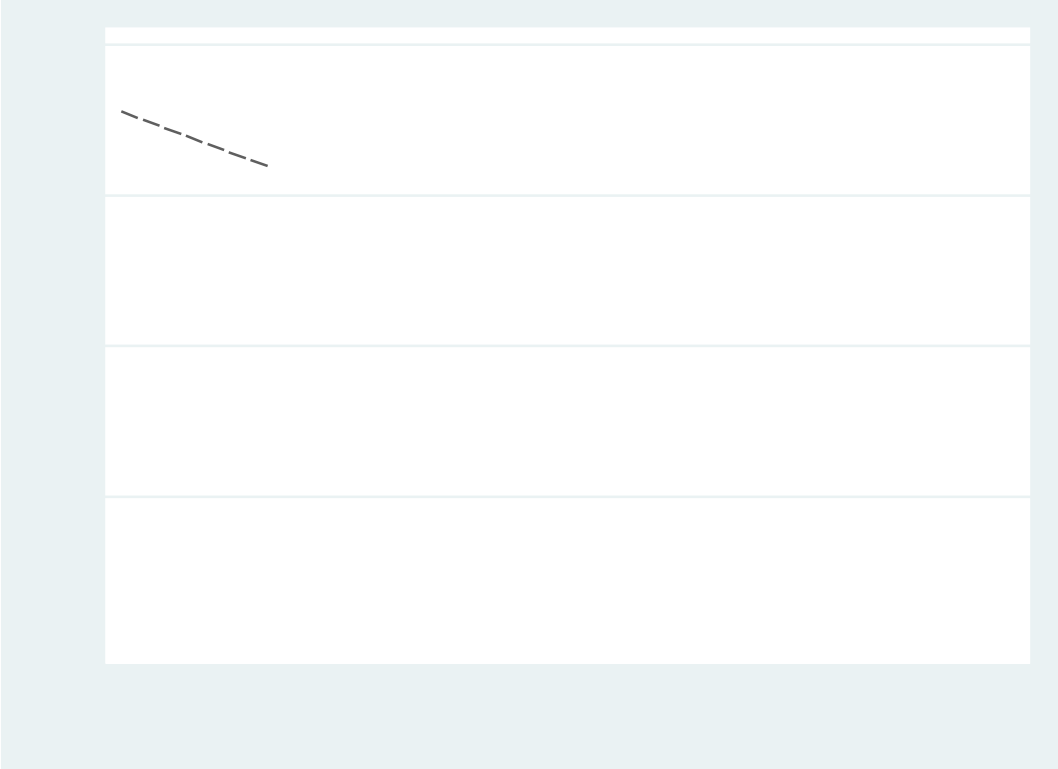
Regime T

Figure 1: The Effect of Regime Type on the Probability of PTA Formation



Note: To compute these predicted probabilities, we use the estimates in the first column of Table 2. The continuous variables are set to their medians. Post-Cold War is set to 1 and the remaining dichotomous variables are set to 0.

Figure 2: The Effect of Veto Players on the Probability of PTA Formation



Note: To compute these predicted probabilities, we use the estimates in the first column of Table 2. The continuous variables are set to their medians. Post-Cold War is set to 1 and the remaining dichotomous variables are set to 0.

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