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Veto players and terror

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Abstract

Democratic regimes have been linked to terrorism for contending reasons, with some scholars claiming democracy increases terrorism and others claiming it decreases terror. Corroborating evidence has been used for both relationships leading to the following puzzle: why do some democratic regimes seem to foster terrorism while others do not? We offer an explanation based on Tsbelis's veto players theory. Beginning with the assumption that terror groups want to change government policy, we argue that the more veto players present in a political system, the more likely the system is to experience deadlock. Given the inability of societal actors to change policies through nonviolent and institutional participation, these systems will tend to generate more terror events. We also explore different methods for estimating terrorism models. We identify several ways to match the data with the proper statistical estimator and discuss implications for terrorism research. Finally, we use new data from the Global Terrorism Database (GTD) that was previously unavailable. These data allow us to use different operational definitions of terrorism and to identify homegrown terror events.

Keywords

democracy, global terrorism database, terrorism, veto players

Introduction

The academic study of political terrorism has been stunted for several reasons including the lack of an agreed upon definition of the central concept (Schmid, 1983), the lack of available data on domestic terror events (Lafree & Dugan, 2007), and the relative lack of scholarly attention to the topic before 11 September 2001 (Kennedy & Lum, 2003). Understanding when, where, and why some oppositional groups use terror while others do not is a critical issue for both policymakers and scholars.

Democracies have recently been the primary targets of suicide terror campaigns as insurgent groups have attempted to coerce their stronger, democratic opponents into conceding (Pape, 2003). Many reasons are offered to explain the nexus between democracy and terror. It may be that civil liberties protections in democracy allow for terror groups to mobilize, stockpile weapons, and act (Gurr, 1979). Or constraints on the executive limit the ability of a leader of a democracy to pursue harsh counter-terror policies that might reduce terror activity (Li, 2005). Finally, minority ethnic or religious groups in democracies may pursue terror to change policy that is immovable through formal political channels due to the group's electoral weakness (Ross, 1993).

Although different mechanisms are offered for democracy's effect on the likelihood of experiencing terror (Ross, 1993), the preponderance of empirical evidence demonstrates that democracies encounter terrorism with more frequency than authoritarian regimes (Gurr, 1979; Turk, 1982; Chenoweth, 2010). Eubank & Weinberg (1994, 1998), using a variety of data and methods, consistently find that democracies experience more terror than their authoritarian counterparts. Li & Schaub (2004) also find a positive association

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between terrorism and the presence of democracy. Li (2005) disaggregates democracy and finds that some attributes of democracy, such as constraints on the executive, promote terror, while others, like democratic participation, diminish the frequency of terror attacks. Eyerman (1998: 152), in contrast, finds that well-established democracies reduce terrorism and suggests that, 'democracies discourage terrorist activity by providing non-violent alternatives for pursuing policy changes in the target state'. To round out the possibilities, Piazza (2008) finds an inconsistent relationship between democracy and terror.

Results from these studies suggest that something about democracy might promote terror. These competing theoretical claims and mixed empirical evidence lead us to question why we see terrorism in some democracies but not others. We offer a simple model of terrorist incentives that utilizes Tsebelis's veto players approach to generate a testable hypothesis – as the number of veto players in a political system increases, terrorism becomes more frequent. The results of our model offer an explanation for why terrorism is more likely in certain democratic states. Our findings also potentially explain the results of previous empirical studies. Previous empirical work on the subject has primarily used the ITERATE dataset which relies on transnational terror events, omitting cases of domestic terror. Since most of the arguments about democracy and terror relate to the interaction between state and domestic terror groups, having data on domestic terrorism is likely important to evaluate these arguments. We use a new database that includes both transnational and domestic terror events. These data also allow us to change the operational definition of terror to probe the robustness of the results depending on how we define terrorism. Quantitative approaches to the study of terrorism have mostly adopted an event count approach to model terrorism. While this approach has benefits and is our primary approach, it overlooks the extensive number of country-years where no terrorism is reported. As several previous scholars note, part of the problem in estimating these models is that terror events are sometimes underreported by media, thus inflating the number of zeros in country-level terrorism databases. To explore the robustness of the results, we also estimate the relationship between veto players and terrorism using models that explicitly accommodate an inflated set of zero values. We first use zero-inflated count models and then adjust the dependent variable to estimate logit models, rare event logit models, and Tobit models. As a further measure of robustness, we model different partitions of the GTD data. And finally, we estimate the primary model using ITERATE data to determine if the results hold with the transnational dataset. With the exception of the model using ITERATE data, the substantive findings of this research hold across these various modeling techniques, different operational definitions of terrorism, and several approaches used to limit potential biases in terrorism data, making us more confident in the statistical results.

Democracy, veto players, and terror

Why do some democracies experience terror events while others do not?[1] This puzzle is important because we have conflicting theoretical claims and mixed empirical evidence. Most importantly, the puzzle is relevant because the consequences reverberate for institutional design. Is there something inherent in certain democratic institutional designs that promotes terrorism?

Many democratic designs are possible. The relationship between the legislature or assembly and the executive can vary in such designs as presidential systems, parliamentary systems, and premier-presidential systems (Shugart & Carey, 1992). Shugart & Carey (1992: 7–8) argue that voters have to trade off efficiency, or 'the ability of elections to serve as a means of voters to identify and choose among the competing government options' and representativeness, or 'the ability of elections to articulate and provide voice in the assembly for diverse interests'. While others have made claims about how the relative stability and quality of policy outcomes relates to whether a democracy is a presidential or parliamentary system (Linz, 1994), Shugart & Carey (1992) show that within presidential systems there is a wide variety of institutional designs that affect stability and policy choices. It is unclear how this distinction among these institutional choices may affect terrorism. Where government breakdown is possible, opposition to the government may choose guerrilla war or insurgency against the state rather than terrorism. Terrorism tends to be a tactic used by extremists who lack support, who are weak, and who face a strong opponent (Lake, 2002).

Lijphart (1999: 1) offers a different scheme for dividing up democracies. In contrast to Shugart and Carey and their interest in looking at executive–legislative relations, his claim is that 'clear patterns and regularities

[1] We define terrorism as the threatened or actual use of illegal force and violence to attain a political goal through fear, coercion or intimidation. A more specific list of criteria is provided below.

appear when . . . institutions are examined from the perspective of how *majoritarian* or how *consensual* their rules and practices are'. In short, the majoritarian model suggests the majority of the electorate's preferences should be pursued by the government in power. Systems that implement this model attempt to produce policy outputs that are consistent with majority opinion. In contrast, the consensual model suggests including as many people's preferences as possible in government output. As Lijphart (1999: 2) suggests, the consensual model's 'rules and institutions aim at broad participation in government and broad agreement on the policies that the government should pursue'. Policy outputs, in this model, should reflect a larger portion of the electorate than the majoritarian model. When empirically testing this difference as it relates to political violence, the findings are mixed. Lijphart finds no association after controlling for factors such as population and development, while Li (2005) finds a slight reduction in transnational terror events for a consensual system as compared to majoritarian or mixed systems.

Tsebelis (1995, 2002) offers a useful model that subsumes the different design schemas as discussed by Shugart & Carey (1992), Lijphart (1999), Linz (1994), and others. Tsebelis (2000: 441) eschews creating typologies of democratic institutional differences, and instead 'start[s] from the final policy outcome of the political game'. His claim is that if different aspects of political systems are important, it has to be due to 'the effects that they have on policy outcomes'. Tsebelis (1995: 292) claims that scholars can compare political systems based on their 'capacity for policy change'. This change is most likely to occur in systems characterized by few 'veto players'.

A veto player, for Tsebelis (1995: 293), 'is an individual or collective actor whose agreement is required for a policy decision'. In contrast to the predictions of Lijphart's consensual model, variation exists within consensual systems. For example, one-party dominant parliamentary systems have fewer veto players than coalition governments. Using the veto players framework highlights this difference. Additionally, one-party dominant parliamentary systems also have fewer veto players than most presidential systems. Using any of the different typologies identified above does not explicate these differences and leads to possibly including different designs that have conflicting impacts on generating policy change.

In addition, the more congruent the various veto players' policy positions, the more likely policy can shift from the status quo. The veto players framework stands in contrast to other institutional explanations that create a dichotomy between different democratic institutional designs (Shugart & Carey, 1992; Linz, 1994; Tsebelis, 1995; Lijphart, 1999). When the United States has a Congress that has the same party across the House and Senate and that same party in the Presidency, the number of veto players is small. In comparison, when there is a divided Congress or when Congress is controlled by a different party than the one that controls the Presidency, the number of veto players increases. The design of the system is the same across these situations; the difference lies in the ability to produce policy outcomes. Other typologies generally ignore this distinction and the impact that it has on producing policy change.

Tsebelis's argument has been influential in political science because the veto players framework 'travels' well. His theory is applicable across time and space and is not bound or conditioned by ad hoc assumptions that might limit the scope or domain of the theory. Its relevance here is obvious because of the growing literature that claims terror campaigns are often strategic and used to induce policy change (Kydd & Walter, 2002; Lake, 2002; Pape, 2003, 2005; Enders & Sandler, 2006). As Crenshaw (1998: 55) argues, terrorists' 'dissatisfaction with the policies of the government is extreme . . . [but] it is not the only method of working towards radical goals'. This suggests identifying the conditions that states create that make terrorism a likely choice by radical opponents of the state. Additionally, Crenshaw (1998: 56) notes that, 'the existence of extremism or rebellious potential is necessary to the resort to terrorism but it does not in itself explain it'. For Crenshaw, terrorism is likely to occur when groups are weak vis-à-vis the state, most people disagree with the goals of radical groups, and the groups themselves fail to mobilize support.

In building on work by Crenshaw and Lake, we develop a model of a group's expected use of terror. We offer some simple assumptions derived from both the burgeoning strategic approach and the veto player's logic. First, we assume that dissident groups use terror to change policy. Groups that oppose the state can use a variety of tools to challenge the state. They can lobby, mobilize protests, form insurgent groups, form political parties, and participate in many other ways. Regardless of the form of political participation, we assume that it is instrumental and directed at influencing policy outcomes. Many groups that utilize terror campaigns are seeking, among other policy goals, to establish their own state or at least an autonomous region.[2] Euskadi Ta Askatasuna or ETA, who have primarily targeted Spain, is one such

[2] ETA, IRA, Chechens, Sendero Luminoso, Hamas, and the PLO are notable examples.

organization. The goal for this group is a major change in policy by the Spanish authorities. Additionally, the LTTE, or Tamil Tigers, perpetrated terror attacks and high-profile suicide attacks against the Sri Lankan people and state to coerce the government into changing its policy relating to autonomy for the Tamil people.

Second, using the insights of the strategic approach, we assume that the way these groups choose to mobilize depends on how high they perceive the costs of participation and the likelihood of policy change. Where costs for participation are low and the likelihood of changing policy is high, terrorism is not an effective strategy. For example, by all measures of democracy, Canada is extremely democratic.[3] The costs of political participation in Canada are low, and it has few veto players compared with other countries such as France, which at times had as many as ten veto players in a given year. Costs for participation in the political system are as low in France as in Canada, but they differ based on the number of veto players. Consequently, France averaged six fatal attacks per year from 1970 to 1997 while Canada averaged less than one.[4] Canada, like many states that experience terror, has an active minority group that seeks a major policy change from the state in terms of autonomy. In Canada, however, policy change towards the goals of the autonomy-seeking group is possible. For example, pro-sovereignty forces for the Quebecois were able to persuade the Canadian government to hold a referendum on sovereignty for the region. In 1995, the referendum was held but narrowly failed.[5]

Where costs for participation are low, but the likelihood of changing policy is also low, we expect greater incidences of terror.[6] In short, since policy change is the goal of a terrorist group or groups opposing the state, then we expect there to be terror events where changing policy is difficult. Italy in the early 1980s, for example, had a high number of veto players (between five and seven) and was highly democratic. Because of the actions of the Red Brigades, Italy experienced about ten fatal terror attacks per year from 1980 to 1984.[7]

As identified above, one way to measure the likelihood of policy change is by counting the number of veto players. Governments with few veto players are better able to produce policy change compared with those with many veto players. Based on the above assumptions, we derive the following hypothesis:

> *Hypothesis*: Terrorist events are *more frequent* as the number of veto players present in a political system *increases*.

Alternative explanations

As noted above, the veto players framework claims that distinctions between majoritarian and proportional systems (or any other similar institutional typology) in changing or maintaining policy may be difficult to discern. Using Tsebelis's veto players framework leads to better understanding of conflicting empirical findings. If a proportional system has few veto players or congruence among the multiple veto players, the likelihood of terror attacks should decrease, since policy change is more likely. If a majoritarian system has many checks and balances on policymaking, the likelihood of terror attacks is expected to be higher, since policy change is unlikely. Furthermore, a proportional system should expect a higher probability of terror incidents than a majoritarian system with few partisan or institutional checks on power.

Li (2005) argues that transnational terror is reduced by democratic participation. As people are able to participate in the political system, they rely less on terror tactics. He maintains that proportional systems should see less terror than majoritarian systems as they represent a wide proportion of society. In contrast, democratic constraints increase terror incidents because, unlike non-democratic regimes, democracies are unable to use any means necessary to limit terror (Wilkinson, 2001). Li (2005: 283) argues that 'institutional constraints severely weaken the ability of the democratic government to fight terrorism'. In other words, since the government cannot adequately

3 Throughout most of the sample period, Canada is listed by Freedom House as being one of the 'Free' countries. Its polity score has been 10, or the highest level, throughout the sample and its scalar measure of democracy from Gates et al. (2006) is 0.936 with 1.0 representing a 'full' democracy.

4 An even larger difference exists when comparing transnational attacks using ITERATE data. Canada only averaged 1.4 attacks per year while France averaged about 20.

5 Canada, however, has not been immune to terror attacks. The Front de Liberation du Quebec used terror in an attempt to influence government policy in the late 1960s. In the early 1970s they changed tactics and were absorbed by the nonviolent groups pursuing autonomy. For a thorough discussion of this process see Ross (1995).

6 Our expectation is that places where costs of participation are high should experience less terror by opposition groups. Where opposition groups develop in these states, we should expect that they have more societal support and thus are more likely to use tactics consistent with either insurgency or more conventional warfare which require mass support.

7 ITERATE codes Italy as having over 18 transnational attacks in the same time period.

pursue counterterrorism, incidents of terrorism are more likely to occur.

States with high executive constraints may be more susceptible to terror and this could explain why some democracies promote terror. This alternative hypothesis, proposed by Li (2005), seems plausible but has several important deficiencies. First, Li (2005: 282) suggests that 'a country that experiences terrorist attacks often attempts to prevent future attacks by adopting policies that circumscribe the freedom of terrorists'. If this is true, then after lowering their executive constraints, states that previously experienced high terror levels should experience less terror. The evidence supporting this claim is decidedly mixed. The USA had the highest levels of executive constraints throughout our sample but experienced varying levels of terror attacks.[8] In 1981, for example, the number of fatal terror attacks in the USA increased from three to six, while the constraints on the executive stayed the same. What changed, however, was the number of veto players in the system rising from 4 to 5 (with the election of Ronald Reagan, a Republican president facing a Democratic Congress).

Li (2005: 283) seems to agree with the veto players logic when he claims that, 'policy inaction and political deadlock, induced by institutional checks and balances, will increase the grievances of marginalized groups pushing them toward violence'. But another mechanism that Li (2005: 294) claims is important, that a system with executive constraints 'weaken[s] the government's ability to fight terrorism', is not consistent with the veto players logic. Policy change or movements from the status quo may be difficult, but if the status quo is already to fight terror, then movements towards policies that diverge from this status quo may be equally difficult.

Finally, Li's logic is related to the effects that state institutional configurations have on domestic and *transnational* actors. This implicitly assumes that the causal processes for both types are the same. Rather than make this assumption, we investigate whether there are differences. While we expect that domestic actors are incentivized to use terror because of policy deadlock, it is unclear how this deadlock affects subnational actors from other states. Do transnational terror processes follow the same logic?

Regime type and executive constraints in this scenario *may* not matter for the country receiving terror. Instead, the institutional factors of the host country of the transnational terror group may do more to explain how they mobilize and choose targets than the domestic situation of the targeted state.[9] In Lai's (2007) terms, it is the *swamp* or the environment that allows a terror group to operate that affects a group's ability to mobilize for transnational terror. In other words, Al-Qaeda organized in Afghanistan, Pakistan, and other locations to perpetrate attacks against the USA, Saudi Arabia, Tunisia, and Bali. These countries' executive constraints had little to no impact on the swampy environment that bred this terror group.

We agree with Li (2005) that explaining democracy's association with terror requires attention to specific causal mechanisms. As Munck & Verkuilen (2002) point out, most scholars following Dahl (1971) conceive of the concept of democracy as having two attributes or dimensions: participation and contestation. Different indicators of these attributes can be used to proxy the concepts. For example, the right to vote and whether that vote is free and fair are standard measures of participation. To proxy contestation, indicators such as freedom of the press or the right to form political parties may be used. Before Li, little attention in the quantitative study of terrorism linked these specific mechanisms with their effect on terror and tended to speak about the relationship between democracy writ large and terror (Eubank & Weinberg, 1994). Since the number of veto players in a system is not necessarily about democracy, though democratic states tend to have more veto players, our argument to some extent moves beyond this debate. Instead of focusing on how or whether people contest or participate, we instead place our attention on policy change or lack thereof and its effect on terror.

New data, multiple methods

While Li attempted to answer the puzzle of whether democracy leads to more or fewer terror attacks (Schmid, 1992; Eubank & Weinberg, 1994; Ross, 1993), he used data on *transnational* terror rather than *domestic* terror. Since arguments about democratic constraints, participation, and policy change may relate mostly to domestic issues, data on domestic terror might be necessary to adequately test some of these claims made in the literature. Whether there are different explanations for transnational and domestic terror is not settled in the literature. Li likely understood this potential mismatch between theory and data but could only use the limited data that were available at the time. Since then, Pinkerton Global Intelligence Securities (PGIS), a private security firm, has donated data

[8] The USA is coded 7, which means that the executive is at parity or subordinate to other groups in society.

[9] Where the targeted state and home state for the terror group coincide but some third party national is killed in an attack, the logic may be more likely to hold up.

recording domestic and transnational terror over a 27-year span to researchers at the University of Maryland. These researchers have coded and augmented the data and created a database that is several times larger than ITERATE, now referred to as the Global Terrorism Database (GTD) (Lafree & Dugan, 2007).

Using these data, we evaluate the hypothesis described above using all countries in the international system from 1975 to 1997. Ideally, the time period would be extended in both directions, as the nature of terrorism could be different before 1975 and/or after 1997 (Rapaport, 2004). Given data availability, this study is limited to making inferences regarding this time period.[10]

Research design

To evaluate the hypothesis derived from our model, we use a sample that includes 115 countries from 1975 to 1997. Our unit of observation is the country-year, and we have a time-series cross-sectional data structure.

Data

This research combines data from several sources. As mentioned above, the annual counts of terrorist attacks were calculated from the most recent version of the GTD, one of the most comprehensive terrorism incident databases available (see LaFree & Dugan, 2007, for a thorough description of the data).[11] The original collectors of GTD data, PGIS, trained researchers to identify and record terrorism incidents from a wide variety of sources such as wire services, US State Department reports, other US and foreign government reporting, US and foreign newspapers, PGIS offices around the world, occasional inputs from such special interests as organized political opposition groups, and data furnished by PGIS clients and other individuals in both official and private capacities.

Despite the efforts by PGIS to create a comprehensive accounting of all terrorist attacks, relying on open source data introduces inevitable difficulties. In order for an attack to be included, it must be reported by the media. For example, the data only report one terrorist incident in North Korea over the entire 27 year period – an unlikely conclusion. This issue is especially a concern for the current research, since less democratic nations are going to have more restrictive media coverage. In fact, a closer look at the data reveals that collection from the more autocratic counties relies more on international news sources, such as FBIS, Reuters, the London *Times*, and the *New York Times*; and collection from the more democratic countries also includes a wider array of news sources including local outlets, such as *El Pais* in Spain, *Milliyet* in Turkey, and the *Philippines Daily Inquirer*. However, despite the availability of local sources in democratic countries, collection efforts predominately relied upon international news sources. To the extent that news access by international news sources is related to the level of democracy in a country (and potentially its number of veto players), the findings may be vulnerable to what Drakos & Gofas (2006b) call an 'underreporting bias'. As an effort to reduce this bias, we use four techniques. First, we focus exclusively on fatal attacks. It is more difficult for a regime to hide fatal attacks, and thus it is more likely that these attacks are picked up by domestic and/or foreign media than attacks that destroy property or lead to non-life threatening injuries. Second, we estimate models using a freedom of the press measure. If a free press is driving the result, then including this variable should weaken our inferences regarding the effects of veto players on terror attacks. Third, as a test for robustness, we estimate zero-inflated models to help sort countries that have little probability of experiencing terror from those that are likely underreporting due to a lack of democracy or free press. We discuss this in detail in the section on model robustness below. Fourth, analysis was repeated on only democracies. By excluding non-democracies, we will essentially retest our hypotheses on only those countries with a free press.

Figure 1 presents the trend of terrorism attacks from 1970 through 1997 as reported by the GTD (total, known homegrown with ambiguous, known homegrown, and known foreign attacks) and ITERATE.[12] First, looking exclusively at the GTD, we notice that regardless of how we operationalize the attacks, terrorism

[10] While the GTD data are available beyond 1997, efforts to link the perpetrators to their home state end in 1997. Since it is important in the current research to distinguish those attacks that were initiated by perpetrators from the target state from other attacks, we elect to analyze the data only up to 1997. While the GTD data are available beginning in 1970, the data on veto players are only available after 1975.

[11] The GTD data were downloaded in May 2009.

[12] We use the terms *homegrown* and *foreign* instead of domestic and transnational because we only distinguish these attacks by the perpetrators' countries of origin, ignoring the nationalities of the targets. Known homegrown and known foreign include only those cases where the perpetrators were identified by the news source and the GTD staff were able to identify their country of origin. Known homegrown attacks are those where the perpetrators attacked in their own country. Foreign attacks are those where the perpetrator's attack is in another country. All other attacks are referred to as ambiguous. Ambiguous attacks could be those with no attributed attacker or those whose attacker was too obscure to clearly identify a country of origin.

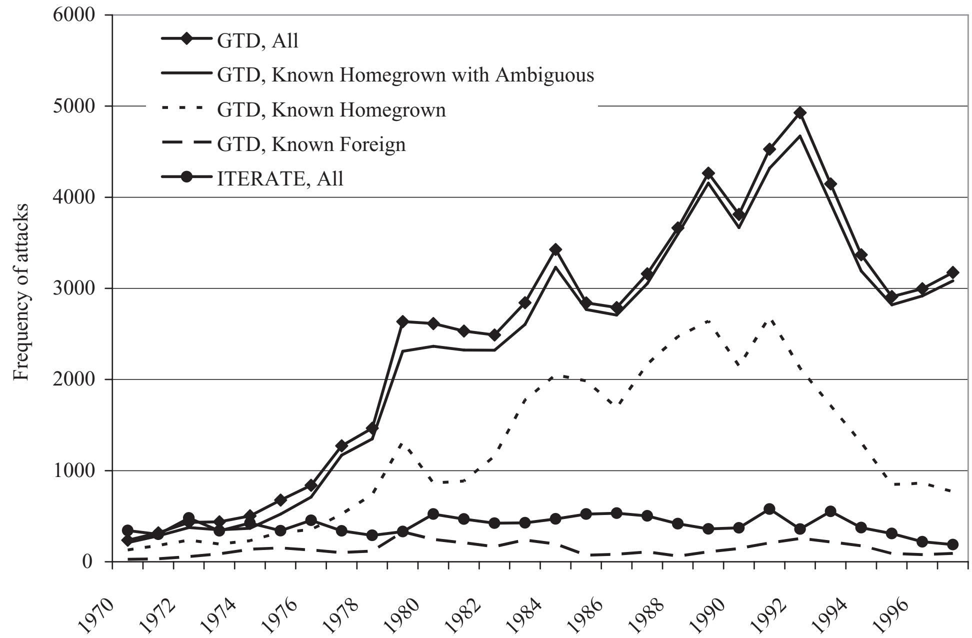

Figure 1. Total terrorism attacks, 1970–1997
Sources: GTD and ITERATE

rose severely from 1970 (237 total, 210 known homegrown with ambiguous, 129 known homegrown, and 27 known foreign attacks) to 1992 (4,928 total, 4,672 known homegrown with ambiguous, 2,124 known homegrown, and 257 known foreign attacks). The largest rise in total activity was between 1978 and 1979, when terrorist attacks rose by close to 80% (nearly doubling). After the global peak in 1992, terrorism begins to drop, ending with more than 3,100 total attacks, 3,000 known homegrown with ambiguous attacks, more than 750 known homegrown attacks, and just under 100 known foreign attacks in 1997.

The GTD records more than 65,000 terrorist attacks over this period (more than 32,000 known homegrown, and more than 3,500 known foreign attacks). This is just under six times the number of terrorist attacks reported in ITERATE over the same period (11,253).[13] As noted above, the main reason the GTD is so much larger than other secondary terrorist incident databases is because it includes information on both domestic and transnational attacks, which is clearly advantageous for this research. Figure 1 shows that if we relied exclusively on the transnational attacks reported in ITERATE as our metric of terrorist violence, we would fail to recognize the large increase in terrorism activity over this period. According to ITERATE, terrorism fluctuated around its mean of 400 attacks, remaining relatively flat over the entire period ($p = 0.79$).

Another reason for the larger number of cases, which is similarly advantageous, is that the GTD is based on a broader definition of terrorism than the one used by most of the other major open source databases. For an attack to be included in the GTD, it must be an intentional act of violence or threat of violence by a non-state actor. In addition, two of the following three criteria also had to be met for inclusion in the GTD: (1) The violent act was aimed at attaining a political, economic, religious, or social goal; (2) The violent act included evidence of an intention to coerce, intimidate, or convey some other message to a larger audience (or audiences) other than the immediate victims; and (3) The violent act was outside the precepts of International Humanitarian Law (START, 2009). Note that this definition differs from other published material using the 1970–2007 GTD data (LaFree & Dugan, 2007). Since those articles were published, the START Center has synthesized the 1970–1997 GTD to match the criteria of the 1998–2007 GTD to form a more consistent series.[14]

[13] Note that we were only able to identify just over 3,500 known foreign attacks compared with the 11,253 identified in ITERATE. This demonstrates the high probability that a portion of the ambiguous attacks include those by foreign perpetrators.

[14] http://www.start.umd.edu/gtd/about/History.aspx.

Dependent variable

Because of the broad-based definition of terrorism used in the GTD, we calculated the dependent variables three different ways (see note 12). We first tabulated all fatal attacks regardless of the target (*All*). This means we included attacks by terror groups against both home states (*Homegrown*), against states that are different than the origin country of the group attacking (*Foreign*), and the ambiguous cases (*Ambiguous*).[15] We then created a separate measure for *Homegrown* attacks, *Homegrown* +←*Ambiguous*, and *Foreign* attacks. The *Ambiguous* attacks are cases where no perpetrators were attributed to the attacks or it was not possible to identify the perpetrators' home state. It is likely that these attacks are perpetrated by a mix of homegrown and foreign agents. The total number of fatal attacks is partitioned this way to account for different levels of possible measurement error. We know that the models for *Homegrown* attacks are measuring only those attacks by known domestic terrorists, with a certain degree of underreporting. The model for *Homegrown* + *Ambiguous* attacks is likely to include all domestic attacks with a certain degree of overreporting. Finally, the *Foreign* attacks model will exclude all attacks by domestic perpetrators. The total number of fatal terrorist attacks for the countries used in this data is 25,441 (23,103 for known homegrown with ambiguous, 13,379 for known homegrown, and 1,184 for known foreign attacks). By comparing estimates across models using each dependent variable, we will be able to determine how sensitive our hypothesis is to the perpetrators' country of origin. In the robustness section, we also estimate models using a dependent variable based on ITERATE data. This measure includes only transnational attacks and includes both fatal and nonfatal attacks.

Estimation strategies

Since each of our four dependent variables is a count of the number of fatal terror attacks in a given country-year – a skewed discrete distribution – standard regression techniques cannot be used. Instead, much of the previous work in this area estimates event count regression models such as a Poisson (PRM) or a Negative Binomial (NBRM). Li (2005) uses a series of NBRM to estimate the effects of democracy on the number of transnational terror events in a country-year. Following Li, we also us NBRM to model the number of fatal terrorist attacks for each country-year from 1975 to 1997.[16] Because observations measured from the same country are inherently dependent and will artificially deflate the standard errors, we cluster the data by country.

Given that count models are the appropriate approach for dealing with terrorist attack data, there are still issues related to reporting bias or when actual attacks deviate from reported attacks (Drakos & Gofas, 2006a,b). Drakos & Gofas (2006b: 81) suggest, 'cases for such deviations are countries that tend to under-report terrorist incidents, because the press is either censored or in some way controlled'. Variation in terrorism reported cross-nationally could be partially related to this bias. Furthermore, the excess number of zeros in the data might be caused by some states censoring the media.

A press freedom measure might explain this deviation (e.g. Li, 2005). Another approach adopted by Drakos & Gofas (2006a) is to use a measure of democracy thought to be highly correlated with a free press. Instead of controlling for this factor, however, they use this measure to predict excess zeros in a Zero Inflated Negative Binomial (ZINB). Thus, while we present our main results using Negative Binomials, we test the robustness of our findings by estimating ZINB models that use a freedom of the press measure in the inflation portion of the model to explicitly account for excess zeros.

While ZINB models are one approach to modeling the problem of excess zeros and underreporting, we try several other estimation techniques to assure that our results are not *model dependent* or dependent upon the assumptions of the statistical model (Ho et al., 2007). We report estimates from logit, rare events logit, and Tobit models to probe the robustness of the results.[17]

Variables

Our central independent variable is the number of veto players in a political system. Two variables commonly used to proxy this concept include the checks variable from the Database of Political Institutions and the *polcon* index from the Political Constraints Dataset.

Checks is a measure of 'the number of decision makers whose agreement is necessary before policies can be changed' (Beck et al., 2001). It is a count of the number of veto

[15] We also looked at the data including and excluding attacks against the military. We include these attacks in our models, but excluding them does not change any of our inferences.

[16] The distribution of values is highly skewed right with zero to five attacks making up 80% of the data.

[17] After estimating a ZINB model, we perform a Vuong test to probe whether using the inflation process is necessary over using a PRM or NBRM.

players, accounting for whether they are independent, and of different parties, adjusted for electoral rules. The more checks in a system, the larger the values of the variable. In our estimation sample, *checks* ranges from 1 to 16 with a mean value of approximately 2.5. *Checks* is coded as 1 for all autocracies. This limits our ability to discuss the role that having more or less veto players within an autocracy has on the likelihood or expected counts of terror. Since we also control for democratic participation and freedom of the press, we can be confident, however, that the measure of differences in veto players is not simply a proxy for differences in regime. To deal with this issue and provide variation within autocracies, we combine the *checks* measure with data from Gandhi & Przeworski (2006). We add a veto player to an autocracy if the autocrat must deal with a legislature. We add another veto player if there are two or more parties in this legislature. Finally, we add a veto player if the dictator is civilian. Gandhi & Przeworski (2006) argue that military dictators should be more able to contain the military, but civilian dictators have to consider the military as a veto player. For an extended discussion of this logic, see Gandhi & Przeworski (2006: 17). We call this new measure *veto*. Returning to our hypothesis, we expect terror events to be more likely as *veto* increases. Additionally, we expect that when at least one attack occurs, increases in *veto* lead to more terror attacks.

An alternative measure of the veto players concept comes from Henisz (2002) and is an index of political constraints called *polcon*. *Polcon* measures the possibility of policy change given the number of independent branches of government with veto power. Low levels of *polcon* represent a system where policy constraints are low, while high levels of *polcon* represent a system where policy constraints are high.[18]

We include a group of control variables to reduce the possibility that the relationship between our key independent variable, *veto*, and terror events is spurious. To measure the degree of democracy in a polity, we use two separate indicators.[19] As mentioned above, we use an indicator to proxy contestation and one to proxy participation. We use the *press freedom* measure used by Li (2005) and developed by Van Belle (1997). This measure is used as an indicator of contestation and as one way to deal with the issue of underreporting. We also use Vanhanen's (2000) modified democratic participation measure used by Li (2005) to proxy *participation*. If claims concerning either dimension of democracy's influence on terrorism supersede our veto players argument, either or both of these measures should 'wash out' the influence that the number of veto players has on the expected counts of terror events within a polity. Many studies of political violence use the Polity scale (Marshall & Jaggers, 2001) as a uni-dimensional indicator. Recent research shows that this may lead to faulty conclusions as the coding rules for Polity 'explicit[ly] reference ... civil war in the coding rules for the variable' (Vreeland, 2008: 402). Moreover, the political participation components of *Polity* are coded in the middle when factionalism occurs, which is a situation where competition between groups is often contentious and violent (Vreeland, 2008). Certainly, terrorism is related to this type of factionalism. The modified Vanhanen measure avoids this problem of relating participation to terrorism *by definition*.

Population, or the natural log of the population, is a common control variable in studies of political violence. Since we are attempting to explain why some countries generate more terror events than others, controlling for population allows us to mitigate the possibility that states with more people simply produce more veto players and are thus likely to have terror events. *Development*, or Gross Domestic Product (GDP), is thought to reduce the likelihood that people in a given polity will have grievances against the state. Where development is high, fewer people should be willing to violently challenge the state. GDP is often used as a measure of development but also proxies for a different concept – the *military capacity* of a state. While it is difficult to disentangle which concept this indicator proxies, it is a standard control in most political violence studies (Fearon & Laitin, 2003; Collier & Hoeffler, 2001). The *region* that a state is part of also may explain some of the reasons that it generates terror events. For example, Latin America in the 1960s and 1970s had a large number of terror events. To control for this possibility, we include a dummy variable for each region of the world from the Correlates of War data.[20] Finally, we include a measure of a state's past history of terror. There are several issues related to dealing with a

[18] The lack of variation within autocracies is a similar problem for *polcon*. Most autocracies have a value of zero. It is not clear how we could combine this index with the autocrat data to improve variation within this subgroup. We do not use *polcon* in the main estimations. The results for this measure have the expected sign but are not significant in many of the estimations. *Polcon* is significant and positive in logits and some other estimations. Given the lack of variation in autocracies and the high correlation with this measure and some of the other independent variables, the results for this measure are much more tentative.

[19] We describe each indicator in greater detail in the online appendix available at http://mypage.siu.edu/jkyoung.

[20] We exclude Latin America from the analyses and use it as the reference category when comparing coefficients across the regions.

Table I. Negative binomial models of fatal terrorism attacks, 1975–1997

Variable	*Model 1 (Homegrown)*	*Model 2 (Homegrown + Ambiguous)*	*Model 3 (Foreign)*	*Model 4 (All)*
Veto	0.202* (0.100)	0.212** (0.062)	0.207* (0.082)	0.221** (0.061)
Free Press	0.022 (0.397)	0.300 (0.309)	0.695 (0.634)	0.429 (0.318)
Participation	0.045** (0.018)	0.021¤ (0.012)	0.013 (0.015)	0.018 (0.011)
Population	0.643** (0.143)	0.526** (0.099)	0.236 (0.092)	0.471** (0.093)
Development	0.048 (0.055)	0.033 (0.036)	0.016 (0.038)	0.042 (0.034)
War	2.664** (0.334)	2.048** (0.267)	0.432 (0.365)	1.970** (0.273)
Cold War	0.654 (0.400)	0.030 (0.213)	0.195 (0.286)	0.098 (0.200)
Durable	0.339* (0.135)	0.369** (0.096)	0.329** (0.114)	0.342** (0.093)
Europe	2.037** (0.709)	1.437* (0.560)	1.025 (0.826)	1.275* (0.537)
Africa	1.764** (0.599)	1.341* (0.434)	1.532* (0.776)	1.445** (0.418)
Asia	2.397** (0.600)	1.502** (0.439)	1.820* (0.780)	1.573** (0.428)
America	0.783 (0.662)	0.515 (0.454)	1.155 (0.805)	0.633 (0.424)
Past Terror	0.112** (0.040)	0.056** (0.020)	0.479 (0.471)	0.054* (0.022)

Robust standard errors clustered on country are in parentheses next to coefficient estimates.
Two-tailed tests, $^{**}p < 0.01$, $^{*}p < 0.05$, ¤$p < 0.10$.

measure of past history. As Brandt et al. (2000) and Brandt & Williams (2001) show, using a lagged dependent variable in event count models can lead to biased and/or inefficient estimates. While Brandt et al. (2000) and Brandt & Williams (2001) offer a couple of solutions depending on the dynamics of the data, neither can be implemented with panel data. We instead follow Li (2005) and estimate an average level of attacks over time (*past terror*) as a proxy for a state's historical experience with terror.[21]

Results

Table I presents the results for four Negative Binomial models – *Homegrown* fatal attacks, *Homegrown+Ambiguous* fatal attacks, *Foreign* fatal attacks, and *All* fatal attacks. The coefficients show how the variables of interest affect the number of terror attacks in a state. Thus, a positive coefficient suggests an increase in the number of terror attacks and a negative coefficient suggests the opposite. Turning to the estimates, we find that the more veto players present in a system, the more fatal attacks expected. The results hold when we include *homegrown, ambiguous, foreign*, and *all* cases. Therefore, we conclude that our hypothesis, which states that increases in the number of veto players in a political system leads to *more* terror events, is supported. Model 1 shows that each additional veto player is expected to increase the number of terror events by 22% (exp(0.202)).

[21] Summary statistics for the estimation are available in the online appendix available at Summary statistics for the estimation are available in the online appendix available at http://www.prio.no/jpr/datasets.asp and at http://mypage.siu.edu/jkyoung.

Turning to the other findings, we see that *free press, population, war, participation*, and *past terror* are positively associated with increases in fatal terror events. These results are generally consistent across estimations. The largest positive effect is the presence of war, which increases the expected number of fatal terrorism events by more than 1,300% in Model 1. While it has a large effect across most models, the result cannot be distinguished from zero in Model 3 or when using data on only known *Foreign* fatal attacks. *Participation* is also insignificant in Model 3 and falls short of the 0.10 significance level in Model 4. Not surprisingly, past experience with terror events increases the number of expected fatal terror attacks across most of the models.

The remaining estimates show a negative association with terrorism. *Cold War*, however, is never significant. *Development* is also consistently negative and also never significant. We find that the durability of the regime decreases the expected count of fatal terror events in all of the models. As the regime and its rules persist, dissidents are less likely to use terror to make policy change. Finally, the negative regional coefficients suggest that most regions produce less terrorism than Latin America. Nearly all of these coefficients are significant suggesting a regional effect.

To probe the robustness of the results, we also estimate a series of models that vary the estimator, the dependent variable, the sample, and the specification to see whether *veto*'s effect is resilient to these changes. Finally, we repeat the NBRM and the ZINB models using ITERATE data to estimate the effects of veto players on total transnational attacks (that include both fatal and nonfatal attacks). Findings from this analysis provide insight into the importance

Table II. The effect of veto on terror attacks given different estimators, coding of the dependent variable, samples, and specifications

Model	Estimator	Dependent variable	Sample/Specification	Veto coefficient (standard error)
Different estimators using GTD data				
5	ZINB	All Fatal Attacks	Full/Only Press Freedom in Inflate Equation	0.221** (0.061)
6	ZINB	Homegrown Fatal Attacks	Full/Only Press Freedom in Inflate Equation	0.194¤ (.105)
7	ZINB	Homegrown + Ambiguous Fatal Attacks	Full/Only Press Freedom in Inflate Equation	0.134** (0.049)
8	ZINB	Foreign Fatal Attacks	Full/Only Press Freedom in Inflate Equation	0.193** (0.073)
9	Logit	Terror Dummy	Full	0.103* (0.049)
10	XTlogit	Terror Dummy	Full	0.118* (0.053)
11	Rare Events Logit	Terror Dummy	Full	0.102* (0.048)
12	Tobit	(Homegrown + Ambiguous Fatal Attacks)/Population	Full	0.197** (0.075)
Different Data/Samples				
13	Negative Binomial	All Fatal Attacks	Only Democracies in GTD	0.068 (0.044)
14	Negative Binomial	ITERATE, All Fatal and Nonfatal Transnational Attacks	Full	0.013 (0.030)
15	ZINB	ITERATE, All Fatal and Nonfatal Transnational Attacks	Full/Only Press Freedom in Inflate Equation	0.013 (0.030)

Robust standard errors clustered on country are in parentheses next to coefficient estimates (except where this option is not allowed – xtlogit and tobit).
Two-tailed tests, $^{**}p < 0.01$, $^{*}p < 0.05$, ¤$p < 0.10$.

of using different measures of terror across multiple sources to better understand its predictors.

Table II displays the results of these tests. First, we used a variety of estimators on the GTD data including ZINB, Tobit, logit, and xtlogit. Across all of the dependent variables, zeros account for a large portion of the observations. Because of this, it is possible that using a zero-inflated model is an appropriate choice. A Vuong test of the ZINB versus a standard negative binomial confirms the notion that a zero-inflated model is preferred ($z = 5.30$, $Pr > z = 0.000$). Other measures of model fit, including the Akaike Information Criteria (AIC) and Bayesian Information Criterion (BIC), also provide evidence that the ZINB is favored over all other count models. When comparing in-sample predictions, the ZINB outperforms the other potential count models. It does especially well at predicting zeros and ones as compared to the ZIP, the PRM, and the NBRM.[22] While there are econometric reasons for using this estimator, as Li (2005: 293) suggests, 'applying the zero-inflated estimator without appropriate substantive theories is problematic'.[23] Since our hypothesis relates to increasing frequency of terror attacks, we lack clear theoretical expectations concerning the factors that might lead to the various types of zeros. Drakos & Gofas (2006b) suggest that the level of democracy or press freedom might be factors that explain the inflation process. Following their approach, we estimate a series of ZINB models (Models 5–8 in Table II) that use the same set of covariates as Models 1 through 4 in the count equation but use the measure of press freedom in the inflation stage.[24] The results are consistent with the previous models.

Another strategy to deal with potential undercounting is to dichotomize the dependent variable. If there is

[22] The results from this test are available in the online appendix available at http://mypage.siu.edu/jkyoung.

[23] This problem is most acute, as Li (2005) notes, because 'the errors of the two equations (probit and negative binomial) are correlated'.

[24] We also tried various measures of democracy and found similar results.

systematic undercounting in some states, dichotomizing could reduce this bias. We created a dichotomous measure where cases with at least one fatal terror attack are coded 1 and 0 otherwise. Using this measure as the dependent variable, we estimated logit, xtlogit, and rare events logit models (Models 9–11). Further, the under-reporting bias could also be treated as a censoring problem. We thus created a fatal terrorism rate per country [(*homegrown*+*ambiguous* attacks/*population* in thousands)*1,000] – or number of attacks per million – and model it using the Tobit (Model 12). The likelihood function of the Tobit combines the likelihood of the probit to predict the probability that a country has at least one terrorist attack with that of the normal to predict the number of attacks. Table II shows that the effect that *veto* has on terrorism is positive and significant in all of these models.

Finally, we estimate the model using different data. First, we limit the GTD data to only democracies (Model 13) as defined by Przeworski et al. (2000). The result is positive; however, the coefficient estimate is about one-third the size of those in Models 4 and 5. This suggests that the role of veto players in non-democracies may also be an important predictor of levels of terrorism. Note that the estimate for Model 13 is also statistically insignificant ($p = 0.12$). We expect that this is due to the substantial loss of statistical power after reducing the number of observations from 2,147 to 742. Finally, we estimate both the Negative Binomial and ZINB models using all transnational terrorist attacks reported by INTERATE (Models 14 and 15). Here we see that *veto*'s effect is much smaller and insignificant in both estimations.

Discussion

The results across the models for the effects of *veto* are generally consistent using different definitions of terror attacks, different model specifications, and different estimators. The important exception is there is a difference when using the ITERATE data. Whether this is due to reporting differences (fatal vs. nonfatal attacks), differences in causal processes between foreign and homegrown attacks, or difference in coding between the different datasets is not entirely clear. In the meantime, we recommend that researchers continue to use different potential measures of terror when conducting cross-national research on terrorism. Additionally, we recommend that scholars continue to probe possible differences between homegrown and foreign attacks.

Some of the inconsistent or inconclusive results for *press freedom, participation*, and *development* may be

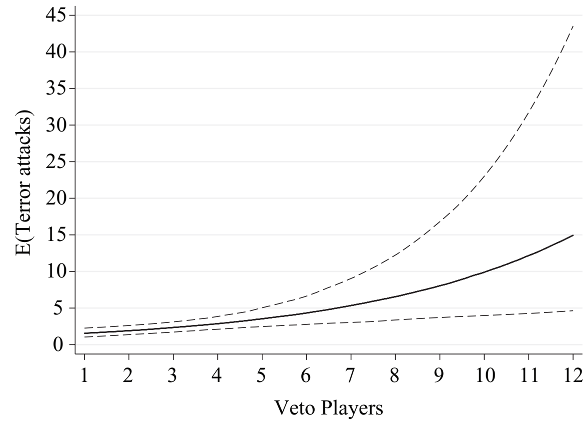

Figure 2. Expected number of terror attacks per country-year given different number of veto players.

due to correlations among these three measures. The three correlate at greater than 0.6 and are clearly competing to explain variance. When using single equation models like the Tobit, the three measures are more consistent but still fail to reach significance in most of the models we estimate. As most other articles acknowledge, this suggests that the relationship between democracy and terror is complicated and that trying to sort out the impacts of the different dimensions is challenging. Given that *veto* holds up across a large portion of the models, this increases our confidence in its impact.

By examining the results of *veto* from Model 1 graphically, we demonstrate how increasing the number of veto players affects the expected number of fatal terror attacks in Figure 2. At the low end, or when a state has a single veto player, the expected number of fatal terror attacks, holding other variables at their mean, is 1.12 plus or minus about 0.5 attacks. When the number of veto players increases to 7, the expected number of fatal terror attacks rises to over 4 attacks with the 95% confidence interval ranging from 2.5 attacks to 6.7 attacks. At the upper end, or when a state has 12 veto players, there is greater variability, but the expected count rises to over 13 attacks.

Implications for future research and policy

Using data from terror events and a variety of modeling strategies that take into account the difficulty in using terror data, we find general support for the role veto players have on the expected number of terror attacks. Our argument and findings suggest that researchers should look beyond the usual institutional typologies of democracies and instead focus on the ability of polities to produce policy change. Since terror groups desire policy change, this focus seems theoretically warranted. Our findings are not conclusive as to whether differences exist between foreign and home-grown terror or domestic, international, or transnational terror. More work should be done theoretically and empirically to draw out potential similarities and differences among these types.

Our findings do demonstrate that the number of veto players in a political system has played an important role in determining a country's vulnerability to terrorist attacks between 1975 and 1997. We are by no means suggesting that countries should change their political structure to accommodate the desires of terrorist organizations. We are suggesting, however, that despite the greater political freedoms in democratic states, increased political stalemate could result in violent unrest. Democracies should be prepared for possible violence when elections produce greater barriers in the policy process.

Despite these conclusions, we acknowledge that this study ends in 1997, raising the question of whether veto players continue to play such an important role today. Since 1997, perceptions of terrorist tactics might have grown to be more severe, due to devastating events of 11 September 2001 in New York and Washington, and more recently in Mumbai. Prior to 2001, terrorism might have been seen as a viable tactic for a broader range of dissidents. Thus, it is important that this model be re-estimated using more recent data that includes terrorism post-9/11. Efforts are currently being made to update the group-level information in the GTD data to the present. We anticipate updating this analysis in the future.

Data replication

The dataset, codebook, log files, and do-files for the empirical analysis are available at http://www.prio.no/jpr/datasets and http://mypage.siu.edu/jkyoung.

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