The Political Effects of Authoritarian Elections*

Mario L. Chacón†

Abstract

This paper investigates the political effects of electoral competition in nondemocracy. Specifically, I explore the effect of political concessions, in the form of multiparty electoral competition, on the likelihood of repression, popular protest, and regime stability. The analysis is guided by a model of nondemocratic politics which incorporates the role of state repression and popular mobilizations as key determinants of regime change. Comparative statics suggests that when authoritarian elections are more competitive, state repression is less likely. The model also shows that for cases in which the opposition is not co-opted by the regime, liberalized autocracies are more likely to democratize. These predictions are then investigated empirically using a panel of countries. The evidence shows that, controlling for time invariant differences as well as global and regional trends in democracy, an increase in the competitiveness of the electoral system decreases the likelihood of repression, has no apparent effect on the incidence of opposition protests, and increases the probability of democracy. These effects are robust and consistent with the motivating theory.

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†New York University Abu Dhabi, PO Box 129188, Abu Dhabi, UAE. Email: mario.chacon@nyu.edu.

1 Introduction

The last quarter of the twentieth century was a period of political liberalization across the world. This transformation included not only an important number of political systems that transitioned from nondemocratic to democratic governments, but also, minor liberalizations in countries that did not become fully democratic. For example, the proportion of regimes having an elected legislature but widely considered to be nondemocratic rose from 58 percent in 1975 to more than 80 percent in 2002 (Keefer 2007). Similarly, the proportion of nondemocratic regimes with more than one autonomous party in the legislature increased from 20 percent in 1972 to 63 percent in 1996 (Gandhi and Przeworski 2006). One of the main results of this global trend is the proliferation of political regimes that combine forms of authoritarianism with seemingly democratic institutions.[1]

An important debate concerns how these liberalizing reforms, which have taken place in nondemocratic societies, influence the political trajectory of these regimes. For instance, what is the relationship between democratic concessions and other strategies of political survival autocrats might use such as state repression? Also, what is the relationship between political openness and popular protest in authoritarian regimes? Lastly, are liberalized regimes on average more likely to democratize than closed autocracies?

To explore these questions, this paper develops a model of nondemocratic politics which incorporates two key elements that have been previously considered in isolation by the literature. First, following recent models of democratization, I model explicitly the use of repression as a substitute strategy aimed to stabilize an authoritarian regime. Second, I include the role of popular mobilization and protest as a necessary condition for political change (e.g., democratization). This framework captures some of the different mechanisms emphasized in the literature (e.g., co-optation of opponents or the advantage of opponents to organize collectively in a liberalized autocracy) and provides sharp predictions about policy determination, repression, and regime change in hybrid regimes.

The theory developed predicts that liberalized autocracies are less likely to rely on repression. This results from the relative value of holding power through repression across different regimes. Since the expected rents from holding office in a liberalized regime are lower than the amount appropriated in a closed regime, the value to an autocrat of holding

[1]See e.g., Diamond 2002 and Levitsky and Way 2002 for an extensive overview of these regimes.

power in a competitive regime is smaller. This implies that an incumbent will be less willing to repress in a liberalized regime relative to a closed regime. The theory also suggests that in cases in which the organizational capacity of the opposition is not undermined by the concessions, liberalized autocracies are more likely to fully democratize. The intuition for this straightforward. If opposition sectors are not "co-opted," democracy is more likely since the incumbent is less likely to repress and opponents are not necessary less likely to demand further concessions.

I empirically explore these predictions using a panel of all nondemocratic regimes in the world during the post-1970 period. Consistent with the motivating theory, I find that "competitive authoritarian" regimes, classified by their level of electoral competition, are less likely to employ state repression. The evidence also suggests that hybrid regimes are not particularly peaceful. In fact, liberalized autocracies are more likely to have insurgencies, strikes, and unrest. Government officials in liberalized autocracies are also more likely to be murdered than their peers in closed regimes. Lastly, the evidence supports the idea that competitive regimes are on average more likely to fully democratize than closed regimes. Overall, the evidence presented casts doubt on the theorized stabilizing effect of democratic concessions under nondemocracy.

This paper contributes to recent works on the role of political institutions under dictatorship (e.g., Gandhi and Przeworski 2006, Myerson 2008, and Boix and Svolik 2013). As mentioned, these works focus on the stabilizing effect of power sharing institutions without considering the long-term dynamics of electoral competition and regime change. Following earlier models on the creation and consolidation of democracy (e.g., North and Weingast 1989, Acemoglu and Robinson 2006), the central element in these formalizations is the capacity of democratic institutions to mitigate commitment problems between autocratic leaders and their supporters (or opponents). Issues such as the conditions under which liberalizing reforms influences the incentives to repress or the potential impact of these on popular unrest and insurrection are largely unexplored in the formal literature.

This paper is also related to an extensive literature on the relationship between political democracy and state repression.[2] Empirical work in this topic has focused mainly on studying the effect of democratic institutions on repressive behavior and human rights abuses. The core finding in most studies is that democratic institutions decrease the likelihood of

[2] See Davenport (2007) for an excellent survey of the theoretical and empirical literature.

state repression (e.g., Davenport 1999; Poe, Tate and Keith 1999; Zanger 2000).[3] This finding, commonly known as the "domestic democratic peace," is depicted in Figure 1., which plots the mean value of systematic purges against political opponents across regimes during the postwar era. It is clear that, on average, democracies are associated with less repressive behavior, particularly during the Cold War period. While this tendency between democracies and nondemocracies has been extensively studied, there is much less research however on how different political institutions influence the occurrence of repression across nondemocratic regimes.

Following the seminar paper of Lipset (1959), a vast number of cross-national studies have focused on the causal relationship between economic growth and democracy (recent examples include: Londregan and Poole 1996, Barro 1999, Przeworski et al. 2000, Boix and Stokes 2003 and Acemoglu e.t al. 2008). The relationship between political institutions and democratization has been less studied, although a couple of recent papers explore how different types of authoritarian regimes are more likely to democratize than others (e.g., Hadenious and Teorell 2007, and Brownlee 2009).[4]

Lastly, the evidence presented departs from the existing empirical literature in two important ways. First, for the sake of parsimony, I formally define only two types of nondemocratic regimes: authoritarian and competitive authoritarian. These types, a stark simplification of the heterogeneity observed across political regimes, highlight one crucial dimension—the electoral competitiveness of the system—which is theoretically relevant for the process of liberalization studied. This approach is different from that taken by recent papers, which employ a broader taxonomy of regime types (e.g., Brownlee 2009).[5] Second, in all models I control for country and time effects in a fully dynamic panel framework,[6]

[3] Although there is a generalized agreement about the negative relation between democracy and state repression, there is less agreement about the model linking these two variables. For instance, some studies claim a non-linear effect, with low and high levels of democracy exhibiting low levels of repression (e.g., Fein 1995), while others find a cut-off effect in which democracy matters but only until a country reaches a certain level of "democraticness" (e.g., Davenport and Armstrong 2004; Bueno de Mesquita et al. 2005).

[4] Other works studying the relationship political institutions and regime change focus instead on movements away from democracy. This literature focuses mainly in the causal link between presidentialism and democratic breakdown (see e.g., Przeworski et.al., 2000 and Boix 2003).

[5] For example, in addition to *electoral authoritarian, hegemonic authoritarian* and *competitive authoritarian,* Brownlee (2009) defines six more regime-year categories (*military, military-personalistic, party hybrid, single-party, personal/military/single-party and monarchy*), based on the work of Geddes (1999).

[6] This empirical strategy has become the norm in the recent literature on democratization and economic conditions (e.g., Bruckner and Ciccone 2011; Boix 2011).

departing from the pooled or cross-sectional models used in literature. This strategy is directly derived from the theory and is useful to deal with unobserved time-invariant variables. Hence, an additional contribution of the paper is to provide a transparent and robust microfoundation to the existing empirical literature on political liberalization and transitions.

2 Theory

Consider a model of political competition between an nondemocratic incumbent labeled I and an opposition group labeled O. The economy consists simply in the production of a natural resource good which generates some exogenous rents labeled R. All rents are collected by the government; thus the incumbent enjoys rents in the amount of R from holding power. Also, suppose I and O are "Downsian" in the sense that they do not have policy preferences and only care about holding office.

There is one large set of citizens (voters), labeled $\mathcal{P}$, with total mass normalized to unity. As in the standard probabilistic voting model (Lindbeck and Weibull, 1987), citizens' preferences consist of two components: economic and ideological. Specifically, all citizens share the same economic preferences, but they may differ in their ideological support for the regime. Formally, the instantaneous utility of $i \in \mathcal{P}$ when group I is in power is given by

$$U_i\left(q, x_I\right) = u(q) + f(|x_I - x_i|) + \xi, \tag{1}$$

where $u(.)$ is an indirect utility function, and $q \in \mathcal{Q} \subset \mathbb{R}^K$ is a vector of economic policies where $K \geq 1$. This utility function is defined over the set of feasible policies, strictly concave, and differentiable. The term x_I represents the ideological bliss point of group I and x_i the ideological bliss point of i. I assume that citizens derive utility from any ideological proximity with the group in power (i.e., f is monotone decreasing). These ideological preferences capture all the non-economic benefits enjoyed by the population from having I in power.[7] Lastly, ξ represents the average (relative) popularity of the regime in the population.

To analyze the effects of electoral concessions under dictatorship influence the path of

[7] This affinity between the citizens and the groups competing for power can be explained in terms of a partisan, ethnic or religious affiliation.

political development, I consider two types of nondemocratic regimes. In the first one, the incumbent is not accountable to the population and the opposition party is banned from competition. This regime represents a "closed" dictatorship in which electoral competition is either not allowed or very low, and is modeled in a simplified way by assuming there are no elections. The second regime considered is a "liberalized" dictatorship regime in which elections take place and both I and O can compete for the popular vote. The main difference between this system and a democracy is that in the former, the incumbent can manipulate the electoral institutions in a way that creates an uneven competition field. The key element in this regime is then that the electoral system installed gives a disproportionate representation to the interest of the supporters of the regime. From now on, this regime -based on lopsided elections- is referred to as "competitive authoritarian" (term coined by Levitsky and Way, 2002). The set of political regimes is then $\mathcal{S} = \{M, C, D\}$, where M stands for closed dictatorship, C for competitive authoritarian, and lastly, D denotes democracy. In the latter, there is free and fair electoral competition and the winner becomes the incumbent for that period.

The dynamics of the game are such that any type of nondemocratic regime, either closed or liberalized, is not stable since the opposition and the voters can organize and demand a full democracy. Following Acemoglu and Robinson (2006), assume that this mobilization capacity is stochastic and fluctuates over time.[8] Explicitly, opponents organize against the regime with probability $p(s)$, and they fail to organize with probability $1 - p(s)$. Cases in which the opposition sectors organize are referred to as "high threat," since during these periods the opposition can disrupt the regular functioning of the economy and demand a political transition.

Lastly, if the opposition solves its collective action problem and organizes, the regime can employ repression and stop the movement.[9] If repression is used, it is always successful, and the incumbent group pays a cost given by

$$\kappa = \overline{\kappa} + \varepsilon, \tag{2}$$

where $\overline{\kappa}$ is an average cost and ε is an iid random variable with a continuous cumulative

[8] The key difference here with the Acemoglu and Robinson model is that the probability of popular unrest is endogenous to the political institutions of the regime.

[9] This aspect of the model can be interpreted as a case in which incumbent is a military group that enjoys the monopoly of violence or as a case in which the military is a perfect agent of group I.

distribution function G defined over $\mathbb{R}_+$. In the empirical analysis of the next section these costs are regime-specific. If repression is not used, the incumbent is forced to introduce democracy. For simplicity, suppose that only I can repress and that the capacity to do so is not dependent upon holding office.

I now explain the payoffs for cases in which elections take place, either in a competitive authoritarian or in a democratic state.

2.1 Electoral Competition

For cases in which elections take place, I and O can credibly commit to any economic policy vector $q \in \mathcal{Q}$, but they cannot make credible commitments regarding their ideological position (represented by x_I and x_O). Therefore, the voters base their decision both on the economic platforms announced and on the known ideological characteristics of these two groups.

Specifically, $i \in \mathcal{P}$ will vote for I when

$$u(q_I) + F(x_i) + \xi > u(q_O), \tag{3}$$

where q_I and q_O are the policy platforms offered by I and O respectively, and

$$F(x_i) \equiv f(|x_I - x_i|) - f(|x_O - x_i|).$$

This last term, which can take negative or positive values, is an individual-specific parameter and it measures $i's$ ideological bias toward the regime. To allow for some degree of heterogeneity in the intensity of ideological leanings, let $F(x_i)$ be an iid random variable uniformly distributed over the closed support $\left[-\frac{1}{2\phi} + \mu, \frac{1}{2\phi} + \mu\right]$, with density $\phi > 0$ and mean $\mu \in (-\frac{1}{2}, \frac{1}{2})$. This distribution, and the voter's policy preferences, are common knowledge to the parties.

Therefore, conditional on the realization of ξ, and given the distributional assumptions, the fraction of citizens supporting a nondemocratic regime ruled by I is

$$\int_{-\Delta u - \xi}^{1/2\phi + \mu} \phi di = \frac{1}{2} + \phi(\mu + \Delta u(q_I, q_O) + \xi), \tag{4}$$

where $\Delta u \equiv u(q_I) - u(q_O)$. This fraction represents the "true" support for I in the population.

Now, the popularity shock ξ is also a random variable, independent and identically distributed, characterized by a uniform distribution with support $\left[-\frac{1}{2}, \frac{1}{2}\right]$.[10] Assuming a majority rule, and given the vector of policies offered by I and O and the ideological distribution in the population, the probability that I wins an election in a democratic state is simply

$$\begin{aligned}\pi^D(q_I, q_O) &= \Pr\left[\frac{1}{2} + \phi\left(\mu + \Delta u(q_I, q_O) + \xi\right) > \frac{1}{2}\right] \\ &= \frac{1}{2} + \mu + \Delta u(q_I, q_O).\end{aligned}$$

Conversely, the electoral probability of the opposition in a democratic state is $1 - \pi^D(q_I, q_O)$.

The electoral competition in a non-democratic state is modeled in a simple way by assuming that parties compete for the votes of only a subset of citizens. Specifically, suppose that a fraction $1 - \lambda$ of citizens do not participate in elections and in return the regime collects a fraction of votes equal to

$$\frac{1}{2} + \phi\left(m + \xi\right), \tag{5}$$

where $m > 0$ is a fixed parameter capturing the capacity of the regime to obtain the support of some citizens by means other than the electoral process (e.g., vote-buying, patronage or coercion). The parameter $\lambda > 0$ measures the importance of elections for the survival of the regime. For $\lambda \to 0$, elections do not play any role in the political system so electoral competition is not a crucial determinant of stability. The popularity shock ξ and the density ϕ are included in (5) to facilitate the comparison with the fully democratic case.

Therefore, the electoral probability for I in a competitive authoritarian election is given by:

$$\begin{aligned}\pi^C(q_I, q_O) &= \Pr\left[\frac{1}{2} + \lambda\phi\left(\mu + \Delta u(q_I, q_O) + \xi\right) + (1-\lambda)\phi\left(m + \xi\right) > \frac{1}{2}\right] \\ &= \frac{1}{2} + \lambda\left(\mu + \Delta u(q_I, q_O)\right) + (1-\lambda)m, \end{aligned} \tag{6}$$

As we see, in a competitive authoritarian state, a fraction of electoral support is invariant to the platforms offered. This captures in a simple way the idea that elections are not a

[10] This shock captures aggregate or "macro" uncertainty during the election at any democratic or non-democratic state.

perfect reflection of voters' preferences in hybrid regimes, and that other exogenous factors (e.g., the effectiveness of patronage or electoral manipulation) account for some portion of the electoral performance.

Clearly, if competitive authoritarian institutions provide some type of electoral advantage to the regime, as the literature suggests, it must be the case that citizens are, on average, less willing to support the regime in a free and fair election relative to the case in which support is obtained using other "non-electoral" strategies. This electoral advantage is guaranteed by the following restriction[11]

Assumption 1: $\frac{1}{2} > m > \mu$.

The characterization of the (subgame perfect) electoral equilibrium in each regime is straightforward. Since parties do not have preferences over policies, the objective function of I is symmetric to that of O, and this will imply policy convergence between them. Also, in a closed regime, elections are irrelevant so parties are indifferent between any policy vector in Q. Then, we have the following result (see the complete characterization in the Appendix)

Lemma 1 *There exists a unique electoral equilibrium in each electoral regime. In this equilibrium, the incumbent and the opposition converge to the same platform. Then, in any subgame perfect equilibrium, the electoral probability for I is*

$$\pi^s(\mathbf{q}^s) = \begin{cases} \frac{1}{2} + \lambda\mu + (1-\lambda)m & \text{for } s = C \\ \frac{1}{2} + \mu & \text{for } s = D \end{cases},$$

where $\mathbf{q}^s = \{q_I^*(s), q_O^*(s)\}$ is pair of subgame perfect policy vectors in regime $s \in \{C, D\}$.

This lemma states that in a competitive authoritarian regime, the incumbent has a higher probability of electoral victory relative the one they would have in democratic state. Also, as $\lambda \to 1$, we approach the case in which elections are free and fair so the probability of victory is approximately the same in both regimes. This simple specification highlights two key sources of power in a "liberalized" autocracy. The first one, captured by μ, is the

[11] This assumption is made only to make the electoral system in the model compatible with the concept of competitive authoritarian elections theorized in the literature (see e.g., Levitsky and Way 2010) and is not important for the main result.

"true" support in the population and the second one, measured by m, is the support gained "artificially" by the electoral system of the regime. The importance of these two depends crucially on λ.

2.2 Repression

I now analyze how electoral concessions might alter the incentives to repress or to democratize fully. In the case of high-threat from the opposition, the decision for the incumbent to repress in state $s \in \{M, C\}$ is given by

$$\max_{\omega \in \{0,1\}} \omega \left(V^I(s) - \kappa \right) + (1 - \omega) V^I(D),$$

where $\omega = 1$ implies repression and $\omega = 0$ implies democratization. In the case in which there is no repression and democratization occurs, I participates in the new democratic system getting the continuation value $V^I(D)$. This specification provides an important link between electoral competition and regime change. Although the role of authoritarian elites in new democracies is explored in some works (see e.g., Grzymala-Busse 2002), it is not considered explicitly by the current formal literature since it is generally assumed that authoritarian elites and ancien régime parties do not have any role in an emerging democracy.[12]

The decision to democratize depends crucially on: the costs of repression, the expected payoff under regime s, and on the continuation value under democracy. For instance, the probability of repression under full authoritarianism is given by

$$\Pr \left[R - \kappa > \pi^D(\mathbf{q}^D) R \right] = G \left[\left(\frac{1}{2} - \mu \right) R - \overline{\kappa} \right], \tag{7}$$

and by

$$\Pr \left[\pi^C(\mathbf{q}^C) R - \kappa > \pi^D(\mathbf{q}^D) R \right] = G \left[(1 - \lambda) (m - \mu) R - \overline{\kappa} \right] \tag{8}$$

under competitive authoritarianism.

Is it clear to see how some of the parameters of the model influence the probability of repression independently of the state. For example, holding all else equal, an increase in R (the rents from holding power) makes repression more likely. This is effect is consistent with

[12] Probably this is explained by the reliance on static models of political competition in which the outcome of a potential regime change is not fully specified (see e.g., Gandhi 2008).

a standard model of rent-seeking and conflict (e.g., Hirshleifer 1991). Not surprisingly, a decrease in $\overline{\kappa}$, the average cost of repression, increases the likelihood of repression. Possibly more interesting is the relationship between regime type and repression. First, notice how the value of holding power in a competitive authoritarian regime is smaller relative to the one in a closed regime. This is simply because the expected rents from holding office in the former are smaller than the amount appropriated in the latter. This implies that, holding all else equal, an incumbent will be less willing to repress, and hence more willing to democratize, if he faces an organized opposition under a liberalized regime relative to the case in which the regime is fully authoritarian. Second, likelihood of repression in a liberalized regime depends negatively on λ and positively on m. This means that as elections become more important for the survival of the regime and the electoral manipulation is lower, repression is less likely. The intuition of this is that as elections become more transparent, the differential value between a competitive authoritarian state and a democracy decreases. Hence, the willingness to repress decreases. Then we have the following lemma.

Lemma 2 *Holding all else constant, as elections under dictatorship become more competitive, state repression is less likely.*

This result, which show how the openness of the regime alters the *willingness* to repress or to tolerate a more democratic system, is empirically explored in the next section.

2.3 Popular Mobilization

The mobilization capacity of the citizens is modeled in a simple way by assuming that the incentives to act collectively against a nondemocratic regime are analogous to the incentives to participate in the electoral process. Moreover, this organizational capacity is endogenous to the electoral institutions of the regime. Explicitly, citizen $i \in \mathcal{P}$ will participate in a rebellion against a nondemocratic regime ruled by I when

$$w - \psi(s) > u(q_I) + F(x_i) + \xi, \tag{9}$$

where w is the expected utility under rebellion, and $\psi(s)$ the cost of supporting the popular mobilization.[13] All other terms are defined as before.

[13] This expected utility can be interpreted as a reduced-form representation of a more strategic process in which citizens take the probability of success in a rebellion and the cost of joining an unsuccessful movement into account. See Chacón, Robinson and Torvik (2011) for a richer model of participation along these lines.

The term $\psi(s)$ captures in a simple way the idea that organizations and institutions could influence the cost of organizing against a non-democratic regime. For instance, opposition parties could provide selective incentives for citizens to participate and solve their collective action dilemma. Since opposition parties are weak (or totally absent) in a closed dictatorship, this could imply that the cost of organizing is lower in a competitive authoritarian regime (i.e., $\psi(C) \leq \psi(M)$). This mechanism is compatible with some of case study and cross-national evidence (e.g., Bratton and van de Walle 1997; Howard and Roesseler 2006) which suggests that elections can provide a "window of opportunity" for an opposition movement.

Even if the cost of organizing is lower in a competitive authoritarian regime, other things may not be the same. In particular, the *opportunity cost* of rebellion is potentially greater in a liberalized regime. To see this, let $\Delta w(s) \equiv w - u(\mathbf{q}^s)$, which is simply the utility differential between the expected value of rebellion and the utility for i in regime $s = M, C$. If competitive authoritarian elections shift the economic policy in an important way, the citizens will derive a higher utility in a competitive authoritarian regime. In that case, we will have $\Delta w(M) > \Delta w(C)$. This means that, all other things being equal, citizens will be more likely to challenge the regime under full authoritarianism. The intuition for this is straightforward: a competitive authoritarian regime is more tolerable for the citizens, so they will be less willing to participate in and pay the cost of an organized rebellion. This way the model provides a simple microfoundation for the "concession effect" of democratic institutions under dictatorship highlighted by works in the literature (e.g., Gandhi and Przeworski 2006).

Now, in order to derive the probability of mobilization under different states, suppose that the threat posed by a popular mobilization is credible only if a critical mass $S > 0$ of citizens are mobilized. If this mass of supporters is not reached, the opposition cannot demand a political transition and the status quo remains unchallenged. Using the same approach as before, the probability of popular mobilization under s is then

$$p(s) = \Pr\left(\xi < \frac{\frac{1}{2} - S}{\phi} + \Delta w(s) - \psi(s) - \mu\right), \tag{10}$$

where ϕ is the density and μ the ideological parameter defined previously.

From (10) is clear to see how S, $\Delta w(s)$, $\psi(s)$, and μ, influence the probability of mobilization. Naturally, the smaller the critical mass of citizens required for a mobilization and the bigger the ideological bias in the population, the more likely it is that the regime will face an organized (credible) opposition. More importantly, the probability of mobilization is endogenous to the electoral system of the regime. For cases in which the net effect of policy concessions is large, mobilization is less likely under competitive authoritarian regime. Otherwise, concessions are not sufficient, and the critical mass of citizens necessary for a mobilization is more likely to emerge in a competitive authoritarian regime. The following lemma presents the exact conditions in which popular mobilization is more likely.

Lemma 3 *The probability of popular mobilization depends on the net effect of policy concessions and on the cost of organizing. Formally, mobilization is more likely under competitive authoritarianism when*

$$\Delta w(C) - \psi(C) > \Delta w(M) - \psi(M),$$

where $\Delta w(s) = w - u(\mathbf{q}^s)$*. Otherwise, mobilization is more likely under a closed system.*

The previous lemma highlights how the probability of popular mobilization in a liberalized regime is higher than in a closed one when the opportunity cost of rebellion under full authoritarianism is high (i.e., citizens enjoy policy benefits in a closed regime). Similarly, holding all else constant, citizens are more likely to challenge a hybrid regime when the costs of organizing are substantially decreased by the electoral concessions of such state (i.e., $\psi(C) \simeq 0$). Hence, there are two contrary effects associated with electoral competition in a nondemocracy. First, there is a potential *concession effect* arising from policy benefits under a liberalized regime. This effect is compatible with the stabilizing effect of democratic institutions under dictatorship explored in the current formal literature (e.g., Magaloni 2008; Gandhi 2008; Boix and Svolik 2010). Second, there is also the potential for an *unraveling effect* in cases when political openness encourages the formation of a critical mass of protestors against the regime. In this case, electoral competition could provide a "window of opportunity" for democratization. If the former effect dominates the latter, we will have $p(M) > p(C) > 0$. This means that an organized opposition capable of threatening the regime is more likely to emerge in a closed autocracy. Alternatively, if $p(M) < p(C) < 1$, concessions are not sufficient and the organizational capacity of the citizens is actually enhanced in a liberalized regime. Lastly, for the special case in which $p(M) = p(C) = p$, the mobilization capacity of the citizens is invariant to the political regime (this is the case considered in the Acemoglu and Robinson's framework).

2.4 Democratization

Under which regime is democratization more likely? The model suggests that state repression is less likely in a liberalized regime. Yet, this does not necessary imply that democracy is more likely to emerge in this state since the organizational capacity of the opposition is endogenous to the regime. Thus, it might be the case that the organizational capacity of the opposition is decreased in a liberalized regime such that overall effect of the likelihood of democracy is negative. If this is the case, seemingly democratic institutions under dictatorship will be associated with greater regime stability. In the opposite case, liberalization increases the chances of a credible mobilization and this can increase the likelihood of democracy. Then, to specify the likelihood of democratization we have to analyze the *joint* probably of repression and unrest.

Formally, the likelihood of democratization in state $s \in \{M, C\}$ is given by

$$p(s)\left[1 - G(\chi(s)R - \overline{\kappa})\right], \tag{11}$$

where

$$\chi(s) = \begin{cases} 1 - \pi^D(\mathbf{q}^D) & \text{for } s = M \\ \pi^C(\mathbf{q}^C) - \pi^D(\mathbf{q}^D) & \text{for } s = C \end{cases}.$$

The likelihood of democracy is simply the joint probability of popular unrest and no state repression. For the case in which the likelihood of protest is invariant to the electoral institutions of the regime (i.e., $p(M) = p(C)$), it is clear to see from (11) that democratization is more likely under a liberalized regime. Why? The intuition for this result is that when the organization capacity of the opposition is not decreased in a liberalized regime -so there is no "co-optation" under competitive authoritarianism-, democracy is more likely since the incumbent is less likely to repress. Then, the likelihood of regime change depends on the relative probability of unrest in different regimes. The following lemma presents a sufficient condition for a co-optation effect in a liberalized regime which make democratization less likely.

Lemma 4 *Democratization is less likely in a liberalized autocracy only when the organizational capacity of the opposition is substantially decreased by the institutions of the regime. Otherwise, a closed regime is more stable. Specifically, democratization is less likely under a competitive authoritarian regime when*

$$p(C) < \widetilde{p},$$

where $\widetilde{p} = \frac{p(M)[1-G(\chi(M)R-\overline{\kappa})]}{1-G(\chi(C)R-\overline{\kappa})}$.

2.5 Summary and Predictions

The model developed in this section studies the joint occurrence of state repression, popular protest and democratization under different types of authoritarian regimes. In particular, I explore how electoral competition under dictatorship shapes the incentives to repress and allow a real democratic system and how this competition influence the strength of the opposition to the regime. The main predictions of this simple framework, which will motivate the empirical analysis in the rest of the paper, are the following:

1. The theory suggest a negative relationship between electoral openness and state repression under dictatorship (Lemma 2).

2. The effect of electoral competition on the likelihood of protest is undetermined. When elections imply substantial policy concessions and do not lower the costs of organization, mobilization is less likely in hybrid regimes (Lemma 3).

3. Lastly, the stabilizing effect of seemingly democratic institutions is conditioned by the effectiveness of the policy concessions associated and by how these influence the cost of organization. If opponents are co-opted by concessions, a popular threat is less likely and hence democratization is less likely to occur in a liberalized regime (Lemma 4).

3 Cross-Country Evidence

I now explore the predictions of the theory using panel data from all nondemocratic regimes over the period 1972-2002. Specifically, I estimate the effects of democratic concessions, in the form of electoral competition, on the likelihood of democratization, state repression and popular unrest.

The empirical strategy followed is based on within-country comparisons across time. In other words, all models are identified from the within-country variation in the data. This strategy is particularly useful to eliminate a potential source of omitted-variable bias caused by any time-invariant factor that might determine the long-run political equilibrium of these countries, for example historical legacies, cultural traditions or geographical characteristics.

3.1 Repression

Some of the variables in the argument in (7) are difficult to measure, and others are unobservable. For instance, we could find a proxy for R (the average rents appropriated by the government), but it would be difficult to calculate the expected value of these rents under different political regimes. Instead, in the econometric analysis, I focus on the following reduced-form model of repression

$$\chi(s_t)R_c - \overline{\kappa}_c = \beta^r C_{ct} + \alpha_c^r + \eta_t^r, \tag{12}$$

where C_{ct} is a variable measuring the presence of a competitive authoritarian regime in country c at date t. β^r measures the effect of competitive authoritarian institutions on the likelihood of repression. According to the theory, $\beta^r < 0$. The average cost $\overline{\kappa}_c$, and all other time-invariant unobserved variables which could potentially influence the likelihood of repression are absorbed by the country-specific effect α_c^r. Similarly, all other temporal factors affecting the occurrence of repression across time are absorbed by the time effect η_t^r .

The econometric model is identified by exploiting the within-country variation in C_{ct}, taking into account global and country-specific trends in repression. This way we can eliminate unobserved time-invariant variables which may be correlated with the costs of repression and the political regime. For example, the unobserved cost $\overline{\kappa}_c$, which influences directly the decision to repress, is incorporated into the model by including country fixed-effects. Also, the inclusion of temporal shocks (i.e., ν_t^r) is crucial given that international factors may be important in explaining both the global trend towards more democratic institutions and toward less tolerance for state repression. This is evident from Figure 1 which shows a sharp decline in repressive tactics across time, both in democratic and nondemocratic states.

3.1.1 Data

The conditional probability derived in (7) is valid only for a nondemocratic regime (the theory does not considers the possibility of repression under democracy). Thus, the first type of data needed for the estimation is a regime classification to limit the estimation to include only nondemocratic regimes. For this purpose, I use the *status* variable from

Freedom House (2008). This variable is a country-year classification that corresponds to a combination of various ratings which indicates the general state of political rights and civil liberties in the political system. I code countries classified as "not free" or "partially free" by Freedom House at any given year as nondemocratic and the ones classified as "free" as democratic.[14]

The second source of data used is the Political Terror Scale (PTS), assembled by Gibney, Cornett and Wood (2010). This 1-5 scale measures levels of state repression, widely conceived as violations of physical and personal rights carried out by state agents. These violations include: extrajudicial killing, disappearances and political imprisonment. The data used to construct the scale for each country comes from the US Department of State's *Country Reports* and from Amnesty International's *Annual Report*.[15] Although this scale is commonly used in the state repression literature (e.g., Davenport and Armstrong 2004), I use it with some caution since it focuses on violations of integrity rights more than it measures general levels of repression. This indicates that the PTS may underpredict the actual level of repression in countries where the state is highly effective in its use of coercion so therefore relatively few explicit acts of violence are sufficient to keep its population repressed.[16]

Variation in types of authoritarian regimes over time (e.g., transitions from a closed to a competitive authoritarian regime) is captured using the *competition* variable from Vanhanen (1997). This is simply the vote share for all the minority parties and independents in parliamentary (or presidential) elections. For cases in which the voting or party information is not available, this variable is calculated as the percentage of seats in parliament held by all minority parties and independents (see Vanhenen 2002). A value of zero corresponds to cases in which the majority party of the system captures all the seats in the legislature and opposition parties are not entitled to compete in elections. This situation approximates a fully closed nondemocratic regime in which electoral competition is either not allowed or

[14] The results are robust to different regime taxonomies such as the ones provided by the Polity 2 Project of Marshall and Jaggers (2004) or the one constructed by Cheibub and Gandhi (2004).

[15] The PTS used combines two scale variables, one from Amnesty International and the other from the US Department of State. Both variables are a 0-5 scale measuring different levels of repression. The variable used is simply an average of these two scales.

[16] Possibly a more important limitation of the PTS data is the bias found by Quian and Yanagizawa (2009). These authors show that the US State Department tended to under-report the human rights violations of countries which were political allies of the United States during the Cold War.

very low.

This empirical approach used to capture differences in political institutions under dictatorship departs from the more qualitative measures used by works in the literature such as Geddes (1999), Gandhi (2008) or Magaloni (2008). Instead of using a binary indicator for the presence of some democratic elements or some subjective assessment about how competitive the regime is, I use Vanhanen's measure since it is more transparent and captures the great heterogeneity in the political institutions of nondemocratic regimes.[17] Also, a competitiveness measure is more compatible with some of the comparative static results derived in the theory. Then, while theoretically I defined two types of nondemocracies, empirically I classify regimes on a continuum according to their electoral competitiveness.

In order to control for time-varying factors that may influence both popular mobilization and the occurrence of state repression, I include log GDP per capita (in 2000 Constant Prices) and the log of total population. These variable are taken from Alan Heston, Robert Summers, and Bettina Atten (2002). In addition, all models include time effects, which are estimated non-parametrically by including year dummies. Lastly, in some specifications I include a full set of country-specific time trends to further control for time-varying factors which may correlated with the political regime. These trends are also included in a flexible way by interacting each country fixed-effect with a linear time variable.

Table 1 contains the descriptive statistics for the variables in the main panel used in the estimation. The sample period is 1976-2006, and each observation corresponds to a country-year unit. In this panel, the mean vote share (or seats) for minor parties in parliament is 17 percent and the mean value of the repression scale equals 2.8. For comparison, I also report the same the descriptive statistics for all democratic countries during the same time period. The average PTS score for democracies in this period is 1.64, and the mean vote share for minority parties and independents is approximately 53 percent.

[17] Some indicators used in the literature do not capture important features highlighted in the theory. For instance, a dummy variable indicating the presence of an elected legislature fails to capture different levels of electoral competition between an authoritarian government and their opposition. Given the great heterogeneity in the institutions of nondemocratic regimes, and the fact that institutions such elections and elected legislatures have become common world-wide, I focus on one dimension of the political environment (i.e., competitiveness). This dimension plays an important role in the model and can be measured in a relatively transparent way.

3.1.2 Results

Table 2 presents the estimates of linear (Columns 1-7) and nonlinear (Columns 8 and 9) models, taking the PTS as the dependent variable and the vote share for all minority parties as the main explanatory variable. All models include a lagged dependent variable to capture the persistence of repression across time and to allow for mean-reverting dynamics. All models also account for period-specific effects. The standard errors reported are robust to arbitrary heteroskedasticity and are clustered at the country level to allow for serial correlation.

Column 1 presents the results of a simple linear model controlling only for country-specific effects. This model is estimated by OLS and includes only the lagged value of repression and the contemporaneous measure of competition as covariates. As we see, the electoral competitiveness measure has an estimated negative effect on the likelihood of repression. The point estimate of -0.28 (which is statistically significant at the 5 percent level) implies that an increase of one standard deviation in this variable decreases the long-run value of repression by 13.6 percentage points [$\simeq 0.2 * 0.28/(1-0.59)$]. To get an idea of the magnitude of this effect, a one-standard-deviation in the PTS of the sample represents about 20 percentage points of the scale. Hence, the magnitude of the effect captured by our political competition variable is substantial.

In column 2, I include log-GDP per capita and log-population as controls. The magnitude and significance of the vote share variable is virtually unchanged by the inclusion of these variables. In column 3, I perform a more demanding test aimed at controlling for time-varying factors by including time trends that are specific to each country. Surprisingly, the coefficient on our vote share variable is even larger in this model, although it is also less precise than before (the estimate is statistically significant only at the 10 percent level). The estimate in this specification implies a decrease of 11.4 percentage points [$\simeq 0.34 * 0.2/(1-0.37)$] in repression from an increase of one-standard deviation in the competitiveness of elections during dictatorship.

As it is well known, the fixed-effects estimates reported so far are biased since the lagged dependent variable is mechanically correlated with the contemporaneous error term. Yet is worth mentioning that this bias may be small since the correlation between the lagged dependent variable and the error term becomes negligibly small as $T \to \infty$. Hence, the annual panel may provide a good a approximation in this case. Still, columns 4-7 explore some corrections to this endogeneity problem. First, following Anderson and Hsiao (1982), I implement a simple transformation and estimate a first-differenced Two Stage Least Squares model in which the second lagged (level) of the dependent variable is used as an instrument for the first-differenced equations.[18] Column 4 and 5 presents the results of this model without and then with controls, respectively. This specification confirms the previous result: electoral competitiveness during dictatorship has a negative, statistically significant effect on the likelihood of repression. In this case, the estimated effect of this variable is larger than before.

Columns 6 and 7 explore the full set of moment conditions of the first-differenced model and use not only the second lag but also all further lags of the dependent variable as instruments. This model is estimated by the Generalized Method of Moments framework proposed by Arellano and Bond (1991). When these moment conditions are valid, this estimator is asymptotically more efficient than the first-differenced IV model of Anderson and Hsiao (see Baltalgi 2008, Ch. 8). In a simple model with no controls, the Arellano and Bond (1991) two-step GMM model yields a lagged repression coefficient of 0.42 and competition effect of -0.49, both statistically significant (column 6). These estimates imply a reduction of 17 percentage points in the long-run level of repression from a one-standard-deviation increase in electoral competition during dictatorship. This effect is also significant and equal to 14 percentage points for the same model with controls (column 7).

The GMM results are compatible with the previous results, but should be taken with some caution. First, we cannot reject the null hypothesis that there is second-order serial correlation in the first-differenced errors. The presence of this kind of serial correlation in the disturbances violates the exclusion restrictions of the model. Besides, both models in Column 6 and 7 reject the null hypothesis in a Sargan test for overidentification. This provides further evidence that some of the moment conditions of this model may not be valid. Therefore, these results should be considered tentative.

As an additional test, I estimate the system-GMM model proposed by Blundell and Bond (1998). This model uses not only lagged levels but also lagged differences as additional instruments. This strategy may improve the precision and the consistency of the Arellano-Bond model (lagged levels may be weak instruments given that both repression

[18] In this specification, political competitiveness is treated as an strictly exogenous variable.

and competitiveness are highly persistent across time). The point estimates and the standard errors of this model are very similar to the GMM estimates presented in columns 6-7. Yet this specification also fails the test of no second-order serial correlation in the residuals and the Sargan overidentification test (these results are not reported)

3.1.3 Robustness

As a further robustness test, I also use a non-linear specification which models the PTS as an ordinal variable representing an ordered "space" of repression. This approach is common in the repression literature (e.g., Davenport 2007a) and in recent papers on the political economy of political violence (e.g., Besley and Persson 2009, 2011). Specifically I estimate an ordered probit model - which is simply a generalization of the standard probit model. This model is estimated by Maximum Likelihood. A clear advantage of this approach is that the PTS may contain only ordinal information, so the actual values of the scale are not relevant (see Davenport (2007) for a critique along these lines). An ordered probit addresses this issue since differences between the actual values of the repression categories are not crucial for the estimation. The standard errors in this model are also robust to heteroskedasticity and are clustered at the country level.

The ordered probit estimates are presented in columns 8 and 9, Table 2. The coefficients of interest are positive and highly significant. The direct interpretation of these is not straightforward since the marginal effect of each independent variable depends on the repression cut points associated with each category. To see this, suppose that the PTS, labeled r, categorizes only three different levels of repression; low ($r = 1$), medium ($r = 2$) and high ($r = 3$). Then, in any ordered probit we will have

$$\begin{aligned}
\Pr(r &= 1|\mathbf{x}) = \boldsymbol{\Phi}(\widetilde{r}_1 - \mathbf{x}\boldsymbol{\beta}) \\
\Pr(r &= 2|\mathbf{x}) = \boldsymbol{\Phi}(\widetilde{r}_2 - \mathbf{x}\boldsymbol{\beta}) - \boldsymbol{\Phi}(\widetilde{r}_1 - \mathbf{x}\boldsymbol{\beta}) \\
\Pr(r &= 3|\mathbf{x}) = \mathbf{1} - \boldsymbol{\Phi}(\widetilde{r}_2 - \mathbf{x}\boldsymbol{\beta}),
\end{aligned} \tag{13}$$

where $\mathbf{x}$ is a vector of covariates, $\boldsymbol{\Phi}$ is the standard normal PDF. $\widetilde{r}_1$ and $\widetilde{r}_2$ are the repression thresholds associated with each (observed) category of repression.[19]

[19] More formally, each repression threshold is given by the following process: suppose a latent repression

The marginal effects of any variable $x_k \in \mathbf{x}$ are then given by

$$
\begin{aligned}
\partial \Pr(r &= 1|.)/\partial x_k = -\beta_k \phi(\widetilde{r}_1 - \mathbf{x}\boldsymbol{\beta}) \\
\partial \Pr(r &= 2|.)/\partial x_k = \beta_k [\phi(\widetilde{r}_1 - \mathbf{x}\boldsymbol{\beta}) - \phi(\widetilde{r}_2 - \mathbf{x}\boldsymbol{\beta})] \\
\partial \Pr(r &= 3|.)/\partial x_k = \beta_k \phi(\widetilde{r}_2 - \mathbf{x}\boldsymbol{\beta}),
\end{aligned}
$$

where ϕ is the standard normal density. It is clear that the response probability for the low and the high category of repression is completely determined by the sign of β_k. For instance, in the simplified model with no controls, the estimated coefficient for our competition variable is -0.6 (se$= 0.25$). This implies that the marginal effect in the low repression category is positive and the one in the high repression category is negative. For the intermediate category, the direction of the effect is ambiguous since it depends not only on the sign of β_k but also on the absolute value of $\widetilde{r}_j - \mathbf{x}\boldsymbol{\beta}$, for $j = 1, 2$. Hence, the marginal effect in the intermediate category depends crucially on the (estimated) cut points $\widetilde{r}_1$ and $\widetilde{r}_2$.

To illustrate more clearly the effect of our main explanatory variable on the likelihood of repression, Figure 2 illustrates the estimated response probabilities using the point estimates in Column 8 and taking all the other independent variables at their sample mean. Consistent with our previous estimates, an increase in electoral competition increases the probability of being in the low categories of the PTS and decreases the probability of being in the high categories. This effect is particularly strong for the "intermediate" values of the PTS (2 to 4) and it vanishes for the extreme values (1 and 5). The difference between marginal effects is explained simply by the shape of the distribution. As we see in Figure 2.3, which plots an Epanechnikov kernel of the PTS used in the estimation of column 8, the mass of the empirical distribution is concentrated in the intermediate values of the scale. Hence, for these intermediate categories, the potential for marginal changes is greater. This implies that as competition increases, holding all else equal, the probability of having a low score in the PTS increases. This is consistent with the estimates of the linear specification discussed above.

---

variable $r^* = \mathbf{x}\mathbf{b} + e$, where $e|\mathbf{x} \sim N(0, 1)$. Then, the data is generated by

$$
\begin{aligned}
r &= 1 \text{ if } r^* < \widetilde{r}_1 \\
r &= 2 \text{ if } \widetilde{r}_1 < r^* < \widetilde{r}_2 \\
r &= 3 \text{ if } r^* > \widetilde{r}_2.
\end{aligned}
$$

3.2 Popular Mobilization

Using the same approach, I now formulate a reduced-form model to approximate the probability of mobilization derived in (10). Specifically, suppose that the argument in this conditional probability is such that for any given country c

$$\frac{1/2 - S_c}{\phi_c} + \Delta w_c(s_t) - \psi_c(s_t) - \mu_c = \beta^u C_{ct} + \alpha_c^u + \eta_t^u, \tag{14}$$

where u denotes "unrest" and β^u measures the effect of competitive authoritarian institutions on the likelihood of popular mobilization. Then, C_{ct} is an indicator for competitive authoritarianism, as defined in (12). The terms α_c^u and η_t^u capture country and time-specific effects in the likelihood of mobilization. The former term captures the effect of all the time-invariant parameters of the model and the latter global shocks influencing popular protest.

As mentioned in Lemma 3, the effect of C_{ct} will depend on the total effect of policy concessions and on the cost of organizing under a competitive authoritarian state. Explicitly, we have that

$$\beta^u \lesseqgtr 0 \Leftrightarrow \Delta w_c(M) - \psi_c(M) \gtreqless \Delta w_c(C) - \psi_c(C)$$

Without having a strong prior about the effectiveness of policy concessions in a liberalized regime or about how these concessions might influence the costs structure of organizing, I estimate the probability of mobilization using a panel of nondemocratic countries between 1972-2006. Since a mobilization against a regime can take various forms, I use a broad range of indicators of social protest aimed at the national government. These include non-violent (e.g., riots) and violent (e.g., armed insurrection and coups) manifestations of organized unrest.

3.2.1 Data

The first source of data comes from Freedom House (2008) and is the same regime classification used in the repression analysis. Using the *status* variable from this publication I restrict the sample in analysis to nondemocratic countries. Following the same strategy used in the repression model, competitive authoritarian institutions are measured by the level of electoral competition. The first indicator is based on the *competition* variable from Vanhanen (1997). As a robustness test, I also use a more qualitative measure of political competitiveness based on the *Polcom* component variable, constructed by Marshall and Jaggers (2004).[20] This variable is a seven point scale measuring the competitiveness of elections and participation.[21] Countries in which opposition parties are repressed and participation is restricted receive low *Polcomp* scores. Countries with institutionalized party competition in the absence of coercion are coded with high scores. To facilitate the interpretation of the results, both measures of political competition are normalized to lie in the zero-one interval.

The second source of data is the domestic conflict event coding from Banks (2008). These data consist of events reported in newspapers, primarily in the *New York Times,* and it includes a broad range of social conflict indicators. The variables used and their definition are:

- *Guerrilla Warfare*: Any armed activity, sabotage, or bombing carried out by independent bands or irregular forces and aimed at the overthrow of the present regime.
- *General Strikes*: Any strike of 1,000 or more workers that involves more than one employer and is aimed at national government policies.
- *Riots*: Any violent demonstration of more than 100 citizens involving the use of force.
- *Anti-government Demonstrations*: Any peaceful public demonstration of at least 100 people for the purpose of displaying their opposition to the government policies.
- *Revolutions*: Any illegal or forced change in the top government, any attempt at such change, or any successful or unsuccessful armed rebellion whose aim is independent from the government.
- *Assassinations*: Any politically motivated murder or attempted murder of a high government official or politician.

Some of these indicators are more compatible than others with the concept of popular mobilization defined in the theory. For instance, armed activities by an irregular army,

[20] This variable allows us not only to test the robustness of the results but also is more in line with previous research on the topic which rely on more qualitative measures (e.g., Wright 2008; Brownlee 2009).

[21] This component variable is based on two measures: the regulation of participation (PARREG) and the competitiveness of participation (PARCOMP). These two variables are classified by their degree of participation and openness (e.g., unregulated, sectarian, restricted or transitional).

riots and anti-government demonstrations are more indicative of a potential threat by an organized opposition than political assassinations or revolutions. This last variable, for example, refers mainly to military interventions, which may or may not be associated with a popular mobilization and social unrest. Still, given their frequency, military coups are relevant, and the scope of the model can easily be expanded to account for military and non-military interventions.

It should be mentioned that because this data set is based on newspaper reports, it is geographically biased (Banks 2005). Given that the main source of information is *The New York Times*, the coverage is naturally biased toward countries and regions that are of greater interest to the United States. Moreover, it is natural to expect that the dissemination of this information depends on the political regime in a way that could bias the estimates. For example, democratic countries may report more incidents of conflict and protest simply because they have free media - although many nondemocratic countries also have partially free or almost free media (Egorov, Guriev and Sonin 2009). Thus, the effect of political openness on conflict is potentially biased by differences in media freedom. For these reasons, these indicators, and the econometric evidence linking political institutions and social unrest, should be taken as tentative and as a suggestion for future research.

For the purpose of the estimation, for each country-year in the sample, I construct a binary indicator for each conflict variable. Then, the dependent variable for each conflict model is a dummy variable taking the value of one if each particular episode is reported in that year and zero otherwise. The sample used in the estimation is a yearly panel covering the period 1972-2006. Table 3 presents the descriptive statistics of the samples used. As we see, the most frequent incidents are "revolutions" and "anti-government demonstrations", which occur in approximately 20 percent of the country-years in the sample. General strikes is the less frequent form of protest, occurring in 5-6 percent of all country-years.

3.2.2 Results

Tables 4 and 5 present the results for the different unrest measures. Each column shows the estimation of a linear probability model taking the binary indicator for each category as a dependent variable and a lagged measure of electoral competitiveness as the main explanatory variable. To capture the temporal persistence of some of these measures and mean-reverting dynamics, all models include a lagged dependent variable. Following equation (14), all the models also include country and year-fixed effects. Most country panels used in the estimation are reasonably large; their minimum average length is 23 years. Thus, I ignore the bias caused by the lagged dependent variable and focus solely on OLS (the fixed-effects OLS estimator becomes consistent as the number of periods gets large). All standard errors reported are robust to arbitrary heteroskedasticity and are clustered at the country level.

Columns 1, 2 and 3 explore the non-violent measures of social unrest, namely, riots, general strikes and anti-government demonstrations. Not surprisingly, these indicators exhibit a positive inertia in the sense that the occurrence of protest today positively influences the probability of protest tomorrow. As the estimates show, lagged protest raises the probability of current protest by approximately 13-15 percent. More importantly, the estimated coefficient of electoral competitiveness is positive in all models, although it is only statistically significant in the strikes model (column 2). For the strikes model, the point estimate is small, and indicates that a one-standard deviation increase in the vote share of the opposition candidates is associated with an *increase* of 2.62 percentage points in the likelihood of a general strike. It is worth mentioning that this is contrary to empirical analysis of Kim and Gandhi (2010) on the relationship between labor concessions under authoritarian rule. These authors find that the presence of legislatures and political parties under dictatorship decreases the number of strikes. Although these authors use the same strikes data from Banks (2008), their identification strategy is different since they use a pooled regression framework. As column 2 in Table 4 shows, the negative effect of democratic institutions on the likelihood of strikes is not robust to the inclusion of country and time effects.[22] Table 5, columns 1-3, replicate these models using the qualitative index *polcomp* from Polity. These estimates are comparable to the previous results, although in none of the three models is the competitiveness effect statistically significant.

The rest of the analysis turns to more radical forms of protest and social unrest: guerrilla warfare, revolutions and politically motivated assassinations. Column 4, Tables 4 and 5, presents the results for the guerrilla warfare indicator. As we see, the coefficient on the measure of competitiveness based on vote shares has a *positive*, statistically significant,

[22] Another important difference between my model and the one of Kim and Gandhi (2010), is the inclusion of a lagged dependent variable. Given the possibility of persistence of strikes and mean reversion dynamics, a dynamic model with a lagged dependent variable is more robust than a pure cross-sectional model.

effect on the likelihood of armed activity by an irregular force against a nondemocratic government. The point estimate in this model suggests that a one-standard deviation increase in the competitiveness of the regime will *increase* the likelihood of guerrilla warfare in the next period by approximately 2 percentage points. This effect is quantitatively small (guerrilla warfare is present in 15 percent of all country-years in the sample) and not robust to the qualitative measure of electoral competition (see column 4).[23]

Columns 5 and 6 report the results for the revolutions indicator and for the political assassinations measure. The effect of electoral competition on the likelihood of a revolution is positive, but again, not statistically significantly different from zero. This result is particularly important since these forced and irregular government changes are the most frequent form of power alternation in nondemocratic regimes (see Hadenius and Teorell, 2007). Thus, if a liberalization is a strategy aimed to facilitate policy concessions and prevent an armed rebellion, the evidence suggests that it is not very effective. On the other hand, the competitiveness effect is positive and significant in the political assassinations model. This result is robust to changes in the measure of competition, although the point estimate when using the *polcomp* index is substantially smaller (see column 6). Therefore, the evidence for a negative relationship between political liberalization and regime stability is null. In fact, there seems to be a robust *positive* relationship between electoral competition under dictatorship and murder of high government officials of the regime.

Lastly, column 7 present the results of a linear model using as dependent variable a dummy which takes the value of one if any of the six protest indicators is one in any given year, and zero otherwise. This variable is labeled "total unrest" and is equal to one in more than 40 percent of all the country-years in the sample. The most common form of protest in this aggregated measure is "revolution" (i.e., coups), which is present in more than 20 percent of all country-years. As we see, the effect of lagged electoral competition is positive but not significant. This is consistent with the overall findings of our protest measures.

Overall, the evidence suggests that the relationship between democratic concessions, in the form of party competition, under dictatorship and popular mobilization is rather weak. The effect of having a liberalized dictatorship as opposed to a closed one is generally

[23]Not surprisingly, we find a strong persistence of guerrilla activity over time. The lagged dependent variable indicates that countries having an armed guerrilla at time *t-1* are almost 50 percent more likely to have an armed insurrection at time t relative to ones that do not.

positive for the different measures, although it is not statistically significant nor robust across specifications.

Taking each measure individually, I find that an increase in electoral competitiveness during dictatorship may actually increase the likelihood of guerrilla warfare. Similarly, government officials and politicians in liberalized autocracies are more prone to be murdered than their peers in closed regimes. Thus, the within-country evidence indicates that institutions providing more representation and participation of the opposition may actually hinder the survival of nondemocratic incumbents.

3.3 Democratization

The econometric model of democratization is somewhat more complicated than the previous models since the dependent variable indicates a transition between different states. A common approach in the literature is to use a dynamic, non-linear model (e.g., probit), imposing a first-order Markov process between democratic and nondemocratic states. One problem with such a specification is the difficulty to include country-specific effects due to the incidental parameters problem (Heckman 1981). One solution to this limitation is to specify a parametric distribution for the unobserved fixed effects à la Chamberlain (1980) and estimate the model by Maximum Likelihood. Instead, the main transition model used is a linear probability model estimated by OLS. This model is more robust since it deals with unobserved time-invariant effects in a non-parametric way.[24]

Specifically, I estimate a linear model of the form

$$d_{ct} = \beta^d C_{ct-1} + \alpha_c^d + \eta_t^d + u_{ct}, \tag{15}$$

where d_{ct} is an indicator function taking the value of one if country c is classified as an autocracy in time $t-1$ and as a democracy in time t, and a value of zero otherwise. Following the same notation as before, C_{ct-1} is a lagged variable indicating the presence of a competitive authoritarian regime at $t-1$, so the coefficient β^d will be the main parameter of interest. α_c^d is a country-specific term capturing all the unobserved time-invariant parameters of the model. η_t^d is a time effect capturing common shocks in the

[24]This LPM is similar to the transition model used by Brückner and Ciccone (2010), although these authors focus on the relationship between economic shocks and political change. This analysis is based on the IV methodology originally proposed by Miguel, Satyanath and Sergenti (2004).

global level of democracy, and lastly, u_{ct} is an error term representing all other unobserved factors that change over time and influence the level of democracy.

Model (15) approximates the probability of transition derived in (11). Even if this specification is not the best description of the response probability of interest, it is a convenient approximation since we can deal with unobserved heterogeneity and time effects in a natural way.[25] In fact, this linear model is often a good approximation of the partial effects on the response probability near the center of the distribution of the covariates (Wooldridge 2002, p. 455).

A common approach in the existing works is to include α_c^d and η_t^d in the error term and estimate a transition model similar to (15) using pooled OLS.[26] The validity of this approach depends crucially on the statistical relationship between α_c^d and C_{ct}. A sufficient condition for consistency in this context is given by

$$cov(\alpha_c^d, C_{ct}) = 0 \text{ for } t = 0, 1, ..T_c,$$

where T_c is the number of periods c appears in the sample. Then, if there is some dependence between α_c^d and C_{ct}, ignoring α_c^d will lead to biased and inconsistent estimates. Since α_c^d captures all the time-invariant parameters of the model, which affect not only the decision to democratize but possibly also the decision to liberalize, this moment condition is tenuous in this context.

To address this issue, I apply a fixed-effects transformation to allow α_c^d to be arbitrarily correlated with C_{ct}. Specifically, I subtract the country-specific mean

$$\frac{1}{T_c}\sum_t d_{ct} = \frac{1}{T_c}\sum_t \beta^d C_{ct-1} + \alpha_c^d + \frac{1}{T_c}\sum_t \eta_t^d + \frac{1}{T_c}\sum_t u_{ct}$$

from (15) to obtain the following estimating equation

$$\widetilde{d}_{ct} = \beta^d \widetilde{C}_{ct-1} + \widetilde{\eta}_t^d + \widetilde{u}_{ct}. \tag{16}$$

[25] A common critique to of this model is that for some values of the explanatory variables the predicted probability will lie outside the unit interval. This can be avoided by using any non-linear specification (e.g. Logit or Probit). As explained, a typical non-linear model cannot accommodate fixed and time effects. Since my empirical strategy depends crucially on fixed-effects, the possibility of having predicted value outside the unit interval becomes a second order concern (see Wooldridge 2002).

[26] In addition to this, the majority of works assume a fully parametric model for the composite error $\alpha_c^d + \eta_t^d + u_{ct} \equiv \epsilon_{ct}$ and estimate (15) by Maximum Likelihood. Since the identification of β^d in this setup comes mainly from the cross sectional variation in the data, these models are analogous to pooled OLS estimation (with the caveat that MLE is less robust since the identification also depends on form of the distribution assumed).

Thus, α_c^d is eliminated in this procedure since it is constant across time. Hence, the model places no restriction on the covariance between α_c^d and C_{ct}. Lastly, since the time shocks $\widetilde{\eta}_t^d$ are unobserved, I estimate these effects in a non-parametric way by including time-specific intercepts.

Equation (16) is estimated by standard OLS.[27] The critical identifying assumption in this framework is that, conditioning on α_c^d, all the components of $\widetilde{u}_{ct}$ are uncorrelated with C_{ct-s} for all $t = 1, 2, ..T$ and $s = 1, 2, ., t$.[28] In the presence of time-varying omitted factors determining the political regime and correlated with C_{tc}, this assumption will not hold, and we have the typical omitted variable bias. In the analysis, I follow a "control" strategy to address this concern including some of the time-varying covariates suggested in the literature. Also, as a robustness check, I include region-specific time trends in some specifications to account for time-varying omitted factors driving both liberalizations and democratizations in a regional manner.

The model is estimated for the baseline period 1972-2002. The baseline panel takes observations every fifth year. As a robustness test, I also construct ten-year and twenty-year panels and estimate the same model. The choice of having a time dimension of five years corresponds to the interpretation given to a generic period in the theoretical framework. Given that we want to capture the long-lasting effect of electoral competition, a time span of five or even ten years seems more appropriate than a time span of one year. Still, even if a five-year or a ten-year specification is preferable, as a robustness test, we also estimate one-year panels - with the caveat that serial correlation could affect the relationship of interest, making inference much more difficult.[29]

3.3.1 Data

For the sake of comparability, the main measure for democratic transitions is constructed from the "Freedom in the World 2008 Survey " produced by Freedom House (2008). Following the same strategy as before, I use the *status* variable, which is available from 1972 to 2008, to identify years of transition between nondemocracy and democracy. The *status*

[27] Since the transformed model is in deviations from country-specific means, this model is equivalent to a pooled regression with country-specific intercepts (i.e., a model including country dummies).

[28] This is the so called "sequential exogeneity" assumption (see Wooldridge, 2003).

[29] Another difficulty in a one-year panel comes from the possibility of a long democratic transition such that the increase in competitiveness is coming from the transition itself. This possibility, the serial correlation issues, and the theoretical motivation, tilt the balance in favor of the 5-year and 10-year panels.

of each country corresponds to the combination of various ratings, and it indicates the general state of political rights and civil liberties in the political system. In the estimation, a democratic transition is coded when a country is labeled as "partly free" or "not free" at time $t-1$ becomes "free" during time t. To check the robustness of the results, a second indicator was created in which a political transition is coded only for countries going from "not free" to "free". According to the methodology of Freedom House, movements from "not free" to "free" correspond to more drastic political reforms compared with movements from "partly free" to "free", which should be interpreted as minor democratic transitions.

To check the consistency and the robustness of the transition years identified by Freedom House, a second classification is used: the "Democracy and Development Extended Data Set" by Alvarez, Cheibub, Limongi, Przeworski (2002). This taxonomy is used extensively by the seminar work of Przeworski et al. (2000). As opposed to the Freedom House taxonomy, this classification is more minimalist in the sense that it focuses on some key requirements regarding the level of political contestation. To construct a second indicator of democratic transitions, I use the regime classification variable *reg* which is coded 1 for years of dictatorship and 0 for democratic years. According to this classification, a country is a dictatorship if the legislature (or the executive) is not elected, if there is no more than one party or if there is no alternation in power (see Przeworski et al. 2000, 14-22). The original classification covers the period 1950-1990 and is updated for the 1990-2002 period by Cheibub and Gandhi (2004).

It is worth mentioning that even though the Przeworski index is based on the type of government and on the level of contestation, and the Freedom House classification is based on political and civil liberty characteristics, the timing and the number of transitions are quite similar and consistent. For the period 1972-2002, Przeworski and his coauthors report a total of 72 democratic transitions. Meanwhile, the Freedom House system reports 76 transitions over the same time span (see Table 6 for details).

Following the same strategy, competitive authoritarian institutions are measured by the *competition* variable from Vanhanen (1997). As explained, *competition* is equal to the vote share for all the minority parties and independents in parliamentary (or presidential) elections. Then, as in our previous models, low values of *competition* correspond to closed dictatorships that have very little or no electoral competition. Similarly, high values of *competition* represent liberalized regimes where elections take place and are meaningful, but they coexist with authoritarian practices (e.g., state repression and electoral manipulation). As a robustness test, I estimate all model using also the *Polcom* component variable, constructed by Marshall and Jaggers (2004).[30]

Finally, in the main estimation I include only two control variables, real GDP per capita (in 2000 Constant Prices) and total population. These variable are taken from Alan Heston, Robert Summers, and Bettina Atten (2002). The effect of income per capita is of particular interest given the importance it has received in the literature. We include total population in the model since we want to control for changes in demographics that could have an effect on the over-time variation of democracy. Table 7 summarizes the descriptive statistics of the panels used in the baseline estimation.

3.3.2 Results

Tables 7 and 8 present the results for the base line period 1972-2002. Table 7 uses the democratic transitions indicator according to the regime classification of Freedom House and Table 8 according to the taxonomy proposed by Przeworski and his coauthors. The standard errors in all models are robust to arbitrary heteroskedasticity and are clustered at the country level.

The first column in these tables uses the political competitiveness index from Polity and the second column the percentage of votes (seats) in parliament or congress held by all minorities parties and independents from Vanhanen. These models demonstrate that lagged competitiveness has a positive, statistically significant, effect on the likelihood of democratization. This result is also robust using different regime classifications. The point estimate in column 1, Table 7, implies that holding everything else constant, a one standard deviation rise in this index increases the chances of democratization in the next five-year period by approximately (0.2*0.37)*100≈7,5 percentage points. Similarly, using the percentage of votes for minority parties as a proxy for competitiveness gives a slightly bigger effect of (0.289*0.37)*100≈10,7 percentage points (column 2). Given that democratic transitions are rare events, these effects are also significant from a quantitative perspective. The models in columns 3 and 4 control for income per capita and population size. The inclusion of these variables does not effect the magnitude or significance of the result. This implies

[30] This variable allows us not only to test the robustness of the results but also is more in line with previous research on the topic which rely on more qualitative measures (e.g., Wright 2008; Brownlee 2009).

that the effect of liberalizations is robust to some time-varying characteristics such as economic performance. Lagged GDP per capita has a positive effect on democratic transitions even though this effect is neither robust nor statistically significant. This is consistent with recent empirical findings about the relationship between income and democracy (Acemoglu et al. 2008, 2009). Lastly, population has a negative and significant effect. This relationship is interesting on its own, and to the best of my knowledge has not been explored in the literature.

The remaining columns explore how the result depends on the frequency of the data. Columns 5 and 6 show that the result holds if we take annual observations instead of 5-year observations. The marginal effect of lagged competition in these models is naturally smaller because a transition in this specification is even more rare than in a 5-year panel. Also, the estimates are more precise since the number of observations is inflated by using a higher frequency in the data. To address concerns of serial correlation, this model with annual data is also estimated including the five lags of all the explanatory variables (column 7). I report the p-value of an F-test of joint significance of these lagged variables. The five-year lags of the competitiveness index are significant at standard levels only for the Freedom House classification (the same is true for the vote share measure, but the result is not reported).

Columns 8 and 9 explore the robustness of the results using a lower frequency in the data (i.e., 10-year panel). The results for this panel are similar and consistent with the ones obtained using the 5-year panel specification. The positive and significant effect of competitiveness has a clear interpretation in this case. Given that we have only three decades in the 10-year specification, the model shows that dictatorships that had a high level of political competition during the 1980's (relative to their own level during the 1970's and to the 1980's world-average) were more likely to democratize during the 1990's. Using the Freedom House data, the size of this effect implies that an increase of one-standard deviation in the *Polcomp* index increases the probability of democracy in the next decade by 9.9 ($\simeq 0.32 * 0.31$) percentage points. For the Przeworski et al. (2000) classification, the magnitude of this effect is 12 percentage points (column 8).

Overall, these results show a large, significant, and positive effect of political liberalization under autocracy on the likelihood of democracy. This effect is robust to different measures of electoral competition and to different frequencies in the data.

3.3.3 Robustness

In order to investigate the robustness of the results, I now estimate the model using additional controls, different time periods, and different samples. To explore the dependence of the model to the time period, I first estimate the model taking the full sample of non-democracies for the entire post-war period using the regime data from Cheibub and Ghandi (2004). The results show that lagged competitiveness, measured by the *polcom* index, has a positive and significant effect, comparable and consistent with the previous estimates (Table 9). The effect is robust to the inclusion of income per capita and (log) population. Given that the time period is longer, I also estimate a panel taking observations every 20 years (see columns 5 and 6). In this panel, almost 50 percent of the variation in the transition function is explained by lagged competitiveness and by the country and time effects. The point estimate suggests that a rise of one standard deviation in the competitiveness measure increase the probability of a democratic transition in the next 20-year period by 25.7 percentage points.

The last two columns in Table 9 present the results of the model when using the more recent data on political regimes from Freedom House (2010). This panel includes all the democratic transitions occurring between 2002 and 2008; thus it covers the entire 1972-2008 period.[31] The resulting estimate for the effect of electoral competition is again positive and significant. Interestingly in this sample, GDP per capita is positive and significant, although the effect is quantitatively small. The effect of population on democratization is negative and consistent with previous models.

As mentioned, the critical identifying assumption in the model is that, conditional on fixed country-specific unobservables, the level of competitiveness is uncorrelated with all the unobserved, time varying, factors affecting the decision to democratize. This condition is clearly violated in the presence of time varying omitted variables correlated with u_{ct}. Then, I now expand the model and include a richer set of time-varying covariates which may serve as better proxies for these omitted variables.

The literature on democratization has emphasized the relationship between economic development and democracy. Along similar lines, some works have emphasized the positive

[31] The Freedom in the World survey codes 6 transitions from "partially free" regimes to "free" regimes during the 2003-2008 period. These country-year transitions are: Trinidad-2005, Guyana-2006, Argentina-2003, Ukraine-2005, and Indonesia-2005.

effect of high levels of education on democracy (e.g., Barro 1999, Przeworski et al. 2000, and Glaeser et al. 2004). Although income per capita is usually a good proxy for human capital, some resource-rich countries have low and medium levels of educational attainment and this may lead to an omitted variable bias in the case in which education is a determinant of democracy.[32] Thus, I augment the baseline model and include two direct measures of education: primary and secondary enrollment. These variables are taken from Banks (2008).

Demographic characteristics, which have been largely unexplored in the theoretical literature, may also affect the likelihood of democracy. For instance, the capacity to organize and protest against a nondemocratic regime may be a function of networks and organizations on the ground, and these may be a function of demographics. To account for this, I include a direct measure of urbanization and a measure of population density. The former is calculated as the number of inhabitants per square mile and the latter as the percentage of national population living in cities of more than 100,000. These variables are also taken from Banks (2008).

Columns 2 and 5, Table 10, present an augmented model which includes the full set of economic and demographic controls described above. As shown, these additional controls have no effect on the size and on the significance of our previous results. Moreover, the estimated effect of lagged competition is bigger than in previous analyses. The point estimate for the model using the *polcomp* index is 0.28 (column 2) and 0.33 for the model using the vote share for minority parties (column 5). Although these regressions contain fewer observations (the education data is only available for a restricted set of years), the estimates are precise and significant at standard levels. Table 10 also reports the p-value of a joint significance test for the two education variables. The results suggest that levels of educational enrollment are not an important determinant of democratization.[33] Similarly, the demographic variables in this model are not statistically significantly different from zero.

[32] Empirically, there is little relationship between political competition in nondemocracies and primary education. For the sample used (5-year panel from 1972 to 2002), the cross-sectional correlation between our vote share measure of competition and primary enrollment is 0.09. On the other hand, the correlation between competitiveness and secondary enrollment is 0.369.

[33] This is consistent with the evidence of Acemoglu et al. (2005), which shows that there is no significant relationship between education (measured by the average years of schooling in the total population of age 25 and above) and levels of democracy in a fixed-effects panel framework.

As a further robustness test, columns 3 and 6 include a full set of region-specific time trends in addition to the year effects included before. These regional trends are estimated in a non-parametric way interacting regional dummies with a linear year variable. The regions used are: Latin America, Sub-Saharan Africa, North Africa and the Middle East, Eastern Europe, and East Asia and the Pacific. As we see, the positive effect of lagged competitiveness is robust to the inclusion of region-specific time effects. An F test of joint significance suggests that dynamics specific to each region are important in explaining movements towards democracy. This is consistent with qualitative works analyzing "waves" and "domino effects" of political change (Huntington 1991), and with the more quantitative literature on the spatial and temporal clustering of democratic institutions (e.g., Gleditsch and Ward 2006). Yet, the results show that the democratizing effect of electoral competition is robust to these regional trends.

Lastly, Table 11 presents a sensibility analysis using various subsamples. An important concern in the estimation is that a particular set of countries could be driving the results, undermining the external validity of the model. To explore this possibility, I use regional subsamples and estimate the same model of democratization used previously. Column 1 replicates the result for the entire sample (which excludes all OECD countries), and columns 2-5 uses regional subsamples for Sub-Saharan Africa, Latin America, the ex-Soviet countries and North Africa and the Middle East, respectively. In this sensitivity analysis, the effect of lagged competition is somewhat smaller when we exclude Latin America and bigger if we exclude Sub-Saharan Africa. However, this effect is positive and significant in all the different subsamples. When we exclude the Soviet bloc and the North African and Middle East countries, the estimate is nearly identical to that of the baseline sample. Thus, as one would expect, the results are not driven by these countries or by regional dynamics of democratization.

3.4 Summary

Overall, the reduced form models estimated in this section are consistent with the comparative statics of the theoretical framework. In particular, these models provide an important piece of evidence for the positive effect of electoral competition under dictatorship on the likelihood of a full-scale democratization down the road. Overall, the estimates suggest that a rise in the vote share of opposition parties increases the likelihood of democratization in the next panel period. This effect is significant and quantitatively large. For the five-year specification, the lower bound of this effect is approximately 9% and the upper bound is 13%. The significance and magnitude of these estimates is robust to different frequencies in the data, different subsamples, and to the inclusion of several time varying covariates and regional trends in democracy.

4 Conclusion

Building on a simple theoretical framework, in this paper I derive a series of reduced form models to study the empirical relationship between competitive authoritarian institutions and various political outcomes. These models are then fitted to cross-sectional time series data on repression, democratization, and organized protest in nondemocratic regimes from 1972 to 2006. The main results of this analysis is that competitive authoritarian regimes are less likely to employ state repression, not necessarily more consensual than their closed counterparts, and lastly, more likely to become democracies. Overall, these findings are robust and consistent with the predictions of the motivating theory.

The econometric evidence in this paper is not immune to the standard shortcomings of any observational study. Although all models account for time-invariant country-specific factors affecting the level of democracy, repression, and protest, they are not a perfect substitute for a more design-based procedure (e.g., an instrumental variables identification strategy). Therefore, some of the moments identifying the effect of political liberalization are questionable and open to debate. Still, the conditional correlations uncovered are interesting on their own and consistent with a simple theory of political institutions under dictatorship and regime change.

Some of the results in this paper are complementary to existing works on the literature - although the theoretical interpretation offered here is different. For instance, the negative effect of competitive authoritarian institutions on the likelihood of repression seems compatible with some of the mechanisms proposed in the state repression literature, such as the possibility of electoral punishment by broad constituencies in more democratic regimes (Davenport 2007b).[34] Similarly, the existing cross-country evidence on political institu-

[34] In the framework of this paper, the effectiveness of electoral punishments as a deterrence mechanism would depend on how responsive is the government to the voter's preferences during elections. If electoral manipulation is high, electoral punishments will be a less important factor influencing the use of repression.

tions in nondemocracies and regime change is compatible with the democratization model estimated here. In these analyses, limited multiparty systems during dictatorship are associated with democratization (Hadenius and Teorell 2007, Brownlee 2009) and have no apparent relationship with the stability of the regime (Gandhi 2008, Brownlee 2009). Hence, a central contribution of this paper is to interpret these separate findings in a single framework.

While the evidence is suggestive, more empirical work on the topic is necessary. In this regard, studies moving beyond cross-national comparisons and focusing on subnational variation seem particularly promising. Also, existing cross-country works should be complemented with more "micro" studies that look directly at causal mechanisms explaining the persistence of authoritarian institutions and other forms of political power. An example of such research is the interesting study of Ziblatt (2008), which explores how high concentration of land ownership and electoral motivations shaped the incentives of legislators in pre-1914 Prussia to reform the highly inegalitarian three-class voting system of the time. In a similar manner, Jha (2010) provides micro evidence from seventeenth century England to show how different endowments (domestic vs. overseas enterprises) influenced the decision of members of the Long Parliament (1640-1660) to initiate a seizure of the Crown's executive authority at the time. This type of investigations exploring specific micro-foundations offer many potential venues for research on the topic of political reform and the persistence of authoritarianism.

5 Appendix

5.1 Proof of Lemma 1

Taking $q_O(s)$ as given, the problem for I in an election type s is simply:

$$\underset{q\in\mathcal{Q}}{Max}\left[\pi^s(q, q_O(s))R\right].$$

Using the derivation of $\pi^s(.)$ (in text) and given that u is differentiable, the first order condition of this problem is:

$$\frac{\partial u(q_I^*(s))}{\partial q} = 0.$$

Similarly, taking $q_I(s)$ as given, the problem for O is given by

$$\underset{q\in\mathcal{Q}}{Max}\left[(1-\pi^s(q_I(s), q))R\right],$$

which has the following first order condition

$$\frac{\partial u(q_O^*(s))}{\partial q} = 0.$$

As shown, the first order conditions for I and O are exactly the same. Given the strict concavity of u, the solution is unique so $q_I^*(s) = q_O^*(s)$ which implies that $\Delta u(q_I^*(s), q_O^*(s)) = 0$. Replacing we get the equilibrium electoral probabilities in the text.

5.2 Proof of Lemma 2

In a competitive autoritarian state, the probability of repression G is such that

$$\frac{\partial G}{\partial \lambda} < 0, \ \frac{\partial G}{\partial m} > 0, \ \frac{\partial G}{\partial \mu} < 0, \text{ and } \frac{\partial G}{\partial R} > 0.$$

Hence an increase in λ (importance of popular support in determining the outcome of elections) and a decrease in m (the level of "fake" support) makes repression less likely.

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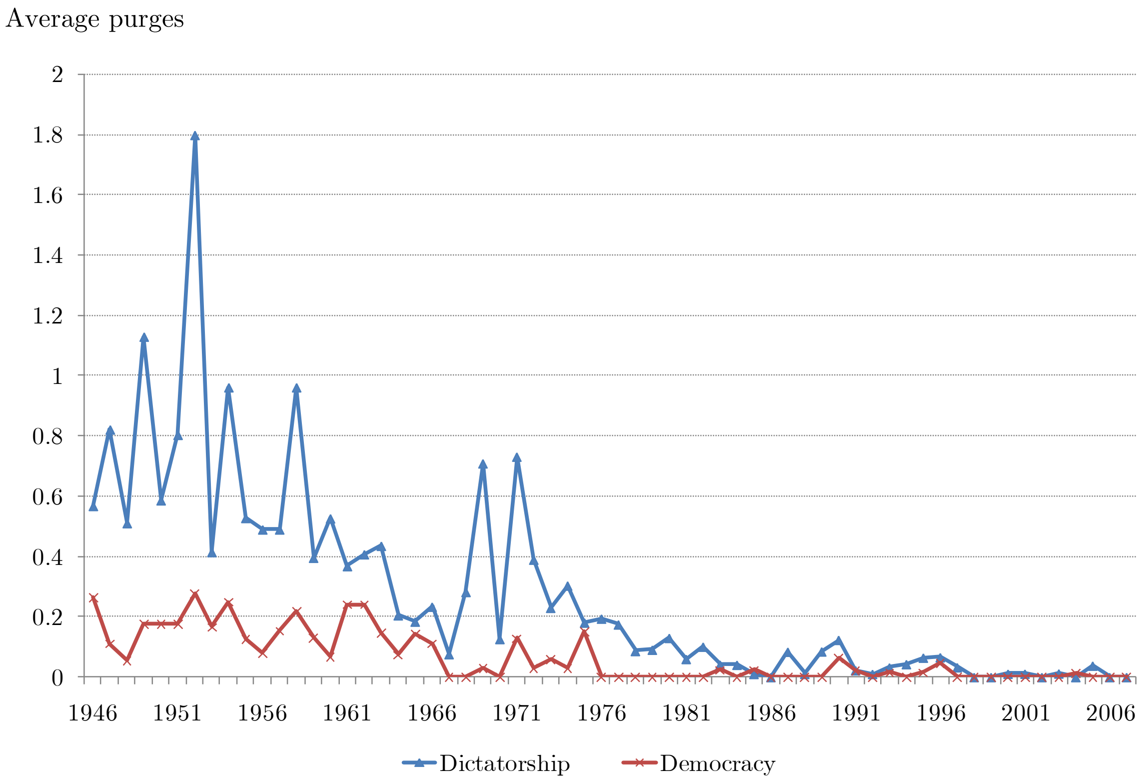
Figure 1
Average Repression by Regime, 1946-2006



Source: Banks (2008) and Marshall and Jaggers (2004). Purges are defined as "any systematic elimination by jailing or execution of political opposition within the ranks of the regime or the opposition". Regimes are classified as democracy if the country reports a positive Polity Score in any given year; otherwise they are classified as dictatorship.

Table 1
Descriptive Statistics
Political Repression, 1976-2006

	Non-democratic Regimes	Democratic Regimes
	(1)	(2)
Repression (PTS)	2.87	1.64
	(1.01)	(0.82)
Percentage votes (seats) for minorities parties$_{t\text{-}1}$	0.17	0.53
	(0.21)	(0.11)
Countries	145	95
Observations	3028	1679

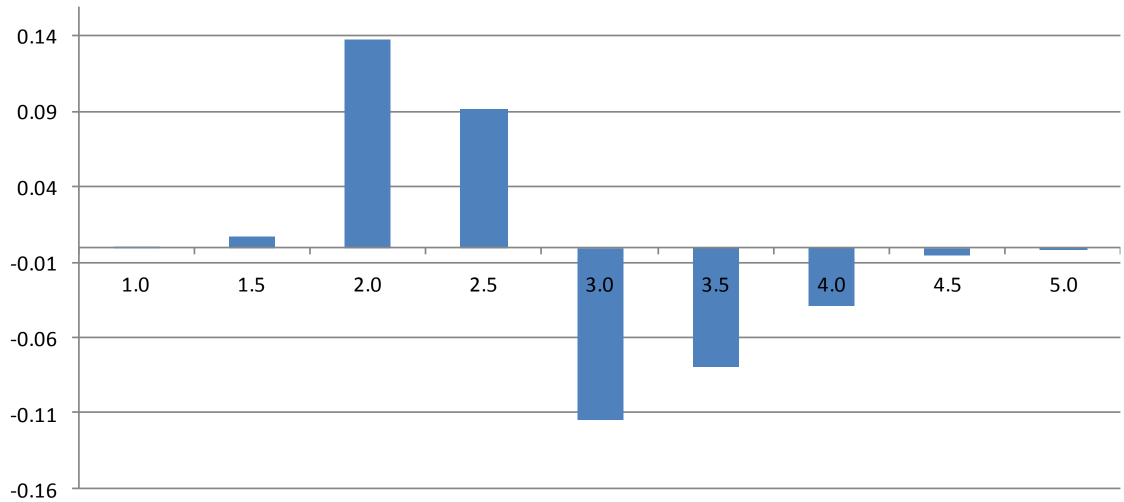
Notes: The numbers reported are averages during sample period. Standard errors are reported in parenthesis. Both columns report the values for the entire 1976-2006 period. The Political Terror Scale (PTS) is taken from Gibney, Cornett and Wood (2010) and the vote share for minorities from Vanhanen (1997). Regimes are classified according to the status variable from Freedom House (2008).

Table 2
Political Repression and Electoral Competition, 1976-2006

Dependent Variable:	Political Repression in Nondemocratic Regimes (1976-2006)								
Model:	Fixed-Effects OLS			Anderson-Hsiao IV		Arrellano-Bond GMM		Ordered Probit	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Repression$_{t\text{-}1}$	0.593	0.571	0.373	0.119	0.147	0.428	0.426	1.226	1.157
	(0.024)	(0.027)	(0.03)	(0.069)	(0.079)	(0.036)	(0.038)	(0.052)	(0.057)
% votes (seats) for minority parties	-0.281	-0.286	-0.344	-0.452	-0.370	-0.497	-0.403	-0.600	-0.607
	(0.123)	(0.146)	(0.19)	(0.137)	(0.157)	(0.219)	(0.267)	(0.255)	(0.296)
Controls		Yes	Yes		Yes		Yes		Yes
Country-Specific Trends			Yes						
Observations	2886	2388	2388	2480	2024	2742	2255	2886	2388
No. Countries	143	133	133	141	128	141	128	143	133
R-squared (within)	0.404	0.405	0.496	0.151	0.143				
Pseudo R-squared								0.344	0.342

Notes: all the standard errors reported in parenthesis are robust to arbitrary heteroskedasticity and are clustered at the country level. All models include year effects. The dependent variable in all models is the PTS from Cornett and Wood (2010). The main explanatory variable is the vote share for all the minority parties and independents in parliamentary (or presidential) elections from Vanhanen (1997). Columns 1, 2, and 3 estimate a linear model with country fixed-effects. Columns 4 and 5 estimate the model using the instrumental variable method of Anderson and Hsiao (1982). Columns 6 and 7 use the Generalized Method of Moments model of Arellano and Bond (1991). These last two models use a double lag of the dependent variable as an instrument of the first-difference model. The electoral competition measure is treated as sequentially exogenous in these models. Lastly, columns 8 and 9 use an ordered probit using the same PTS information.

Figure 2
Marginal Effect of Electoral Competitiveness, PTS



Note: The y-axis represents the estimated response probabilities using the point estimates in Column 8, Table 2.2 and taking all the other independent variables at their sample mean.

Figure 3
Empirical Density PTS
Nondemocratic Regimes, 1976-2006

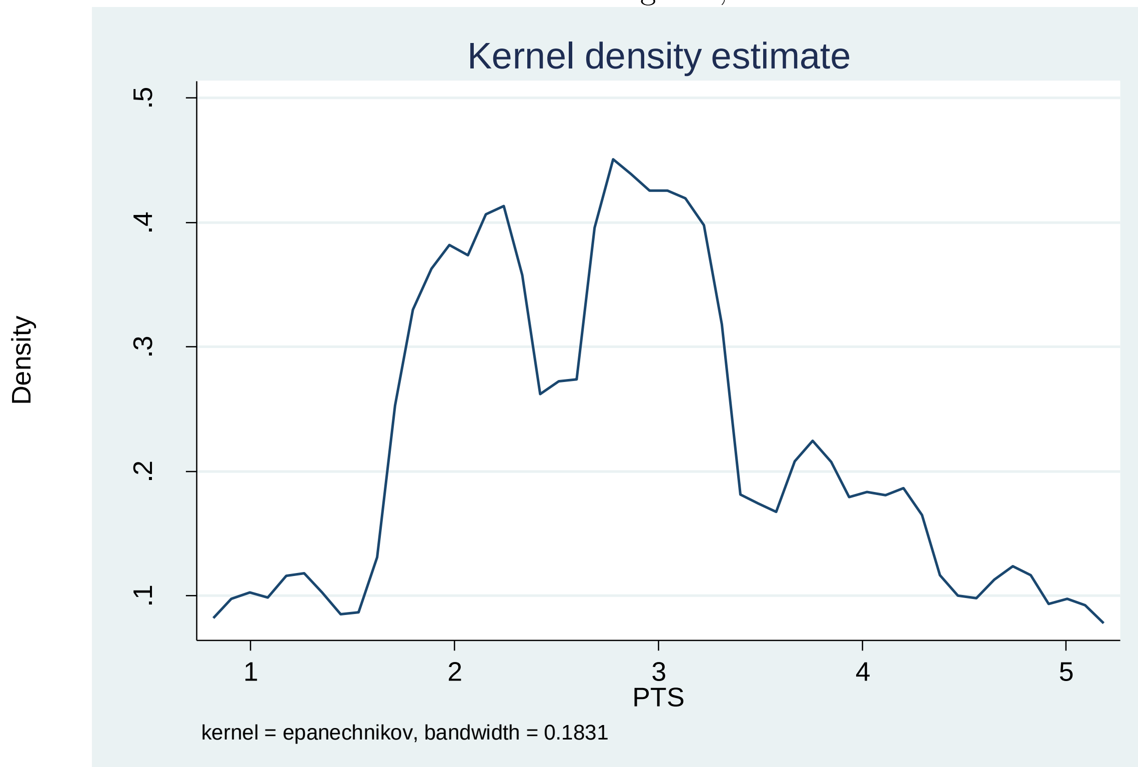


Table S
Descriptive Statistics
Social Protest and Unrest under Dictatorship, 1972-2006

	Panel A	Panel B
	(1)	(2)
Riots	0.138	0.131
	(0.34)	(0.33)
General Strikes	0.065	0.058
	(0.24)	(0.23)
Anti-government Demonstrations	0.195	0.189
	(0.39)	(0.39)
Guerrilla Warfare	0.152	0.139
	(0.35)	(0.34)
Revolutions	0.217	0.195
	(0.41)	(0.39)
Political Assassinations	0.103	0.098
	(0.30)	(0.29)
Total Unrest	0.424	0.408
	(0.49)	-0.49
% votes (seats) for minority parties	0.157	
	(0.20)	
Political Competitiveness Index		0.252
		(0.31)

Notes: All the conflict variables are taken from Banks (2008). These variables are dummies indicating the presence of each type of event during any given country-year. The vote share for minority candidates is taken from Vanhanen (1997) and the Political Competitiveness Index is constructed from the Polcomp index from Marshall and Jaggers (2004).

Table T
Social Unrest and Protest under Dictatorship I, 1972-2006

Model:	LMP with country fixed-effects						
Sample:	1972-2006 (annual panels)						
Dependent variable:	Riots	General Strikes	Anti-Government Demostrations	Guerrilla Warfare	Revolutions	Political Assassinations	Total Unrest
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Lagged dependent variable	0.152	0.140	0.150	0.422	0.297	0.150	0.206
	(0.026)	(0.031)	(0.023)	(0.033)	(0.031)	(0.029)	(0.022)
% votes (seats) for minority	0.020	0.128	0.002	0.099	0.061	0.108	0.068
parties$_{t\text{-}1}$	(0.054)	(0.044)	(0.069)	(0.043)	(0.07)	(0.05)	(0.069)
Observations	3615	3615	3613	3615	3613	3615	3619
No. Countries	151	151	151	151	151	151	151
R-squared (within)	0.040	0.052	0.070	0.206	0.112	0.054	0.083

Notes: the standard errors reported in parenthesis are robust to arbitrary heteroskedasticity and are clustered at the country level. All models include year effects. Dependent variable in all models is a dummy indicator taking the value of 1 if during the year any of the conflicts events were reported, zero otherwise. All conflict events are taken from Banks (2008). The percentage of votes (or seats) for the minority parries is taken from Vanhanen (1997).

Table U
Social Unrest and Protest under Dictatorship II, 1972-2006

Model:	LMP with country fixed-effects						
Sample:	1972-2006 (annual panels)						
Dependent variable:	Riots	General Strikes	Anti-Government Demostrations	Guerrilla Warfare	Revolutions	Political Assassinations	Total Unrest
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Lagged dependent variable	0.135	0.144	0.137	0.434	0.285	0.142	0.188
	(0.03)	(0.035)	(0.027)	(0.036)	(0.034)	(0.031)	(0.024)
Political Competitiveness	0.029	0.027	0.001	-0.012	0.005	0.051	0.028
Index$_{t-1}$	(0.039)	(0.023)	(0.046)	(0.033)	(0.046)	(0.029)	(0.05)
Observations	3217	3217	3217	3217	3215	3217	3221
No. Countries	140	140	140	140	140	140	140
R-squared (within)	0.036	0.052	0.070	0.220	0.108	0.056	0.077

Notes: the standard errors reported in parenthesis are robust to arbitrary heteroskedasticity and are clustered at the country level. All models include year effects. Dependent variable in all models is a dummy indicator taking the value of 1 if during the year any of the conflicts events were reported, zero otherwise. All conflict events are taken from Banks (2008). The Political Competitive Index is constructed from the Polity project, Marshall and Jaggers (2004).

Table V
Democratic Transitions, 1972-2000

Period	Freedom House (2004)	Przeworski et al. (2000)
1973-1980	Argentina (73), Botswana (73), Dominican Rep (78), Ecuador (79), Ghana (80) Greece (74) Grenada (77), India (77), Maldives I. (74), Nigeria (79), Papua NG (76), Peru (80), Portugal (76), Spain (77), Sri Lanka (76), Thailand (75), Turkey (74), Burkina F (78).	Argentina (73), Bolivia (79), Brazil (79), Ecuador (79), Ghana (79), Greece (74), Nigeria (79), Peru (80), Portugal (76), Spain (77), Thailand (75), Uganda (80)
1981-1985	Argentina (84), Brazil (85), Grenada (85), Honduras (84), Uruguay (85)	Argentina (83), Bolivia (82), Cyprus (83), El Salvador (84) Grenada (84), Honduras (82), Nicaragua (84), Thailand (83), Turkey (83), Uruguay (85)
1986-1990	Chile (90), Czechoslovakia (90) Gambia (89), Hungary (90), Korea Rep. (88), Malta (87), Namibia (90), Philippines (87), Poland (90), Suriname (88), Thailand (89), Vanuatu (89), W. Samoa (89)	Bulgaria (90), Chile (90), Comoro I (90), Czechoslovakia (90), Guatemala (86), Hungary (90), Korea rep. (88), Panama (89), Philippines (86), Poland (89), Romania (90), Sri Lanka (89), Sudan (86), Suriname (88)
1991-1995	Bangladesh (91), Benin (91), Bulgaria (91), C Verde I (91), Estonia (93), Guyana (93), Latvia (94), Malawi (94), Mali (92), Mali (95), Mongolia (91), Nepal (91), Panama (94), Sao T/P (91), Slovak Rep. (94), S. Africa (94), Zambia (91)	Albania (92), Bangladesh (91), Benin (91), Burundi (93), C. Verde I (91), Central Af. Rep. (93), Congo (92), Ghana (93), Guyana (92), Haiti (94), Lesotho (93), Madagascar (93), Malawi (94), Mali (92), Mongolia (92), Nepal (91), Niger (93), Sao T/P (91), S. Africa (94), Suriname (91), Thailand (92), Zambia (91)
1996-2000	Bolivia (96), Croatia (00), Dominican rep (98), Ecuador (98), el Salvador (97), Fiji (99), Ghana (00), Honduras (97), India (98), Mexico (00), Papua NG (98), Philippines (96), Romania (96), Slovak Rep.(98), Suriname (00), Taiwan (96), Thailand (98), Venezuela (96)	Cote d'I (00), Guinea-B (00), Indonesia (99), Kenya (98), Mexico (00), Moldova (96), Niger (00), Nigeria (99), Senegal (00), Sierra L. (96), Sierra L. (98), Taiwan (96)

Notes: Using the Freedom House taxonomy, a country experiences a democratic transition in years in which the country is classified as "free" in that year conditional on being classified as "not free" or as "partially free" in the previous year. Similarly, using the Przeworski index, the country reports a democratization for any year in which the country is classified as a democracy, conditional on being classified as a dictatorship in the previous year. The exact year in which the transition takes place is reported in parenthesis.

Table W
Descriptive Statistics
Democratic Transition, 1972-2002

	Transitions Freedom House	Transitions PACL (2000)
	(1)	(2)
Period 1972-2002		
Democratic Transition Indicator	0.09	0.09
	(0.28)	(0.28)
Countries	158	146
Observations	649	682
Panel A		
Democratic Transition Indicator	0.09	0.10
	(0.28)	(0.31)
Political Competitiveness Index$_{t\text{-}1}$	0.18	0.14
	(0.28)	(0.25)
Countries	134	124
Observations	565	525
Panel B		
Democratic Transition Indicator	0.09	0.11
	(0.29)	(0.31)
Percentage votes (seats) for	0.12	0.09
minorities parties$_{t\text{-}1}$	(0.19)	(0.16)
Countries	148	131
Observations	636	575
Panel C		
Ln GDP per capita$_{t\text{-}1}$	7.86	7.83
	(1.03)	(1.05)
Ln population$_{t\text{-}1}$	15.60	15.45
	(1.78)	(1.74)
Countries	135	118
Observations	570	515

Notes: The numbers reported are averages during sample period. Standard errors are reported in parenthesis. The first row reports the values for the entire 1972-2002 period. The first column of Panel A refers to the sample used in column 1 of Table 2.5. Similarly, the second column of this panel refers to the sample of column 1 Table 2.6. Similarly, the first column of Panel B refers to the sample used in Table 2.5, column 2, and the second to the sample of Table 2.6, column 2. Panel C, column 1, refers to the sample of Table 2.5, column 3. Panel C, column 2 refers to the sample used in Table 2.6, column 3. The observations refer to the total number of observations and the countries to the total number of countries used in each panel. The democratic transition indicator takes a vale one if a country is classified as a democracy conditional on being classified as an autocracy in the previous period. In column 1 the classification of regimes is taken from Freedom House (2008) and in column 2 from Przeworski et al. (2000). Real GDP per capita (in 2000 Constant Prices) and total population are taken from Heston, Summers, and Atten (2002).

Table X
Transitions to Democracy, Freedom House

Model:	Fixed-effects OLS								
Dependent variable:	Freedom House Transitions to Democracy, 1972-2002								
Panel:	Five-year data				Annual data			Ten-year data	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Political Competitiveness Index$_{t\text{-}1}$	0.203		0.201		0.094		[0.02]	0.329	
	(0.075)		(0.078)		(0.025)			(0.127)	
% votes (seats) for minority parties$_{t\text{-}1}$		0.289		0.255		0.135			0.449
		(0.112)		(0.111)		(0.042)			(0.195)
Ln GDP per capita$_{t\text{-}1}$			0.069	0.048	0.003	-0.006	[0.56]	0.076	0.062
			(0.035)	(0.034)	(0.008)	(0.010)		(0.067)	(0.058)
Ln population$_{t\text{-}1}$			-0.264	-0.315	-0.029	-0.092	[0.45]	-0.421	-0.486
			(0.107)	(0.111)	(0.026)	(0.034)		(0.161)	(0.170)
Observations	565	636	504	570	2336	2702	2083	244	279
No. Countries	134	148	122	135	135	142	122	108	122
R-squared (within)	0.085	0.101	0.103	0.116	0.034	0.033	0.044	0.234	0.227

Notes: the standard errors reported in parenthesis are robust to arbitrary heteroskedasticity and are clustered at the country level. All models include year effects. Dependent variable in all models is a transition function taking the value of 1 if a country is classified as "free" in time t conditional on being classified as "not free" or as "partially free" in period *t-1*. This classification is taken from Freedom House (2008). Sample in 1-4 is an unbalanced panel with observations taken every 5 years, sample in 5-7 is an unbalanced panel with annual observations and sample in 8-9 is an unbalanced panel with data is taken every 10 years. In column 7 the p-value from an F-test of joint significance of all 5 lags is reported in brackets.

Table Y

Transitions to Democracy, Przeworski et al. (2000)

Model:	Fixed-effects OLS								
Dependent variable:	Przeworski et al. (2000) Transitions to Democracy, 1972-2002								
Panel:	Five-year data				Annual data			Ten-year data	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Political Competitiveness Index$_{t-1}$	0.248		0.259		0.073		[0.31]	0.454	
	(0.096)		(0.098)		(0.034)			(0.172)	
% votes (seats) for minority parties$_{t-1}$		0.358		0.37		0.109			0.559
		(0.145)		(0.149)		(0.057)			(0.306)
Ln GDP per capita$_{t-1}$			0.054	0.015	-0.001	-0.006	[0.50]	-0.019	0.004
			(0.048)	(0.044)	(0.011)	(0.012)		(0.099)	(0.074)
Ln population$_{t-1}$			-0.306	-0.284	-0.071	-0.101	[0.14]	-0.486	-0.347
			(0.109)	(0.101)	(0.029)	(0.034)		(0.185)	(0.142)
Observations	525	575	464	515	2242	2517	2008	232	259
No. Countries	124	131	110	118	113	122	109	102	111
R-squared (within)	0.144	0.144	0.145	0.145	0.047	0.046	0.054	0.292	0.241

Notes: the standard errors reported in parenthesis are robust to arbitrary heteroskedasticity and are clustered at the country level. All models include year effects. Dependent variable in all models is a transition function taking the value of 1 if a country is classified as a democracy at time t conditional on being classified as a dictatorship in period t-1. This classification is taken from Preworski et al. (2000). Sample in 1-4 is an unbalanced panel with observations taken every 5 years, sample in 5-7 is an unbalanced panel with annual observations and sample in 8-9 is an unbalanced panel with data is taken every 10 years. In column 7 the p-value from an F-test of joint significance of all 5 lags is reported in brackets.

Table QP
Robustness Test I, Time Period

Model:	Fixed-effects OLS						Fixed-effects OLS	
Dependent Variable:	Przeworski s Transitions						FH s Transitions	
Sample:	1946-2002						1972-2008	
Panel:	Five-year data		Ten-year data		Twenty-year data		Five-year data	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Political Competitiveness Index$_{t-1}$	0.168	0.174	0.29	0.291	0.692	0.541	0.182	0.19
	(0.070)	(0.075)	(0.120)	(0.129)	(0.226)	(0.245)	(0.064)	(0.066)
Ln GDP per capita$_{t-1}$		0.054		-0.004		0.022		0.07
		(0.043)		(0.079)		(0.194)		(0.03)
Ln population$_{t-1}$		-0.305		-0.515		-0.618		-0.26
		(0.105)		(0.183)		(0.610)		(0.101)
Observations	797	606	368	293	157	125	653	587
No. Countries	128	129	124	105	105	91	136	123
R-squared (within)	0.119	0.134	0.261	0.281	0.496	0.506	0.077	0.099

Notes: the standard errors reported in parenthesis are robust to arbitrary heteroskedasticity and are clustered at the country level. All models include year effects. Dependent variable in models (1)-(6) is a transition function taking the value of 1 if a country is classified as a democracy at time t conditional on being classified as a dictatorship in period t-1. This classification is taken from Preworski et al. (2000). Samples in 1-2, 3-4, and 5-6 are unbalanced panels with observations taken every 5-year, 10-year, and 20-year respectively. The dependent variable in models 7-8 is the same transitions function according to Freedom House (2010). Lastly, the samples in these two models are unbalanced panel with observations every 5 years.

Table QQ
Robustness Test II, Additional Controls

Model:	Fixed-effects OLS					
Dependent Variable:	Freedom House Transitions to Democracy					
Sample/Panel:	1972-2002, five-year data					
	(1)	(2)	(3)	(4)	(5)	(6)
Political Competitiveness Index$_{t\text{-}1}$	0.203	0.284	0.292			
	(0.075)	(0.127)	(0.133)			
% votes (seats) for minority parties$_{t\text{-}1}$				0.289	0.338	0.269
				(0.112)	(0.197)	(0.148)
Ln GDP per capita$_{t\text{-}1}$		0.234	0.367		0.191	0.331
		(0.096)	(0.128)		(0.084)	(0.121)
Urbanization$_{t\text{-}1}$		0.124	0.145		0.183	0.153
		(0.212)	(0.225)		(0.193)	(0.21)
Primary and Secondary Enrollment (*F* test)		[0.68]	[0.68]		[0.91]	[0.83]
Ln population$_{t\text{-}1}$		-0.320	0.241		-0.349	0.212
		(0.463)	(0.45)		(0.409)	(0.44)
Ln population density $_{t\text{-}1}$		0.041	-0.386		-0.108	-0.484
		(0.261)	(0.336)		(0.261)	(0.327)
Regional time trends (*F* test)			[0.03]			[0.06]
Observations	565	198	198	636	212	212
No. Countries	134	103	103	148	110	110
R-squared (within)	0.085	0.271	0.343	0.101	0.26	0.329

Notes: the standard errors reported in parenthesis are robust to arbitrary heteroskedasticity and are clustered at the country level. All models include year effects. Dependent variable in all models is a transition function taking the value of 1 if a country is classified as "free" in time *t* conditional on being classified as "not free" or as "partially free" in period *t-1*. This classification is taken from Freedom House (2008). All samples are unbalanced panels with observations taken every 5 years. The *p*-value from an F test of joint significance for all the regional-specific time trends is reported in brackets. The regions used to construct these trends are Latin America, Sub-Saharan Africa, North Africa and Middle East and East Asia.

Table QR

Robustness Test III, Regional Subsamples

Model:	Fixed effects OLS				
Dependent variable:	Freedom House transition to democracy 1972-2002				
Panel:	Five-year panel				
Sample:	all countries	without Sub-Saharan Africa	without Latin America	without ex-Soviet countries	without North Africa & Middle East
	(1)	(2)	(3)	(4)	(5)
Political Competitiveness Index$_{t-1}$	0.203	0.244	0.139	0.214	0.212
	(0.075)	(0.115)	(0.079)	(0.077)	(0.08)
Observations	565	353	490	497	469
No. Countries	134	92	114	108	116
R-squared (within)	0.085	0.103	0.082	0.072	0.099

Notes: the standard errors reported in parenthesis are robust to arbitrary heteroskedasticity and are clustered at the country level. All models include year effects. Dependent variable in models (1)-(5) is a is a transition function taking the value of 1 if a country is classified as “free” in time t conditional on being classified as “not free” or as “partially free” in period *t-1*. This classification is taken from Freedom House. All samples are unbalanced panels with observations taken every five years.