THE LOGIC OF POLITICAL VIOLENCE*

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This article offers a unified approach for studying political violence whether it emerges as repression or civil war. We formulate a model where an incumbent or opposition can use violence to maintain or acquire power to study which political and economic factors drive one-sided or two-sided violence (repression or civil war). The model predicts a hierarchy of violence states from peace via repression to civil war; and suggests a natural empirical approach. Exploiting only within-country variation in the data, we show that violence is associated with shocks that can affect wages and aid. As in the theory, these effects are only present where political institutions are noncohesive. *JEL* Codes: D74, H40, O11.

I. Introduction

Political violence is the hallmark of weakly institutionalized polities. The starkest manifestation of such violence is armed conflict in the form of civil war. Counting all countries and years since 1950, the average yearly prevalence of civil conflict, according to the Armed Conflict Dataset (ACD), is over 10%, with a peak of more than 15% in the early 1990s. The upper left part of Figure I shows the variable trend in the worldwide prevalence of civil war by year. By contrast, the upper right graph plots the prevalence of civil war by country (since 1950 or independence, if later) against GDP per capita in 1980. Clearly, civil wars are disproportionately concentrated in the poor countries of the world. The cumulated death toll of these conflicts exceeds 15 million people (See Lacina and Gleditsch 2005).

A key feature of civil war is two-sided violence between an insurgent and the government. However, many citizens suffer consequences of one-sided political violence, due to government repression through a variety of infringements of human rights. The Banks (2005) data set reports a stark form of repression viz.

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purges—that is, the removal, by jailing or assassination, of opponents considered undesirable by the incumbent government. Since 1950, about 7% of all country-years are associated with purges, in the absence of outright civil war. The lower left graph in

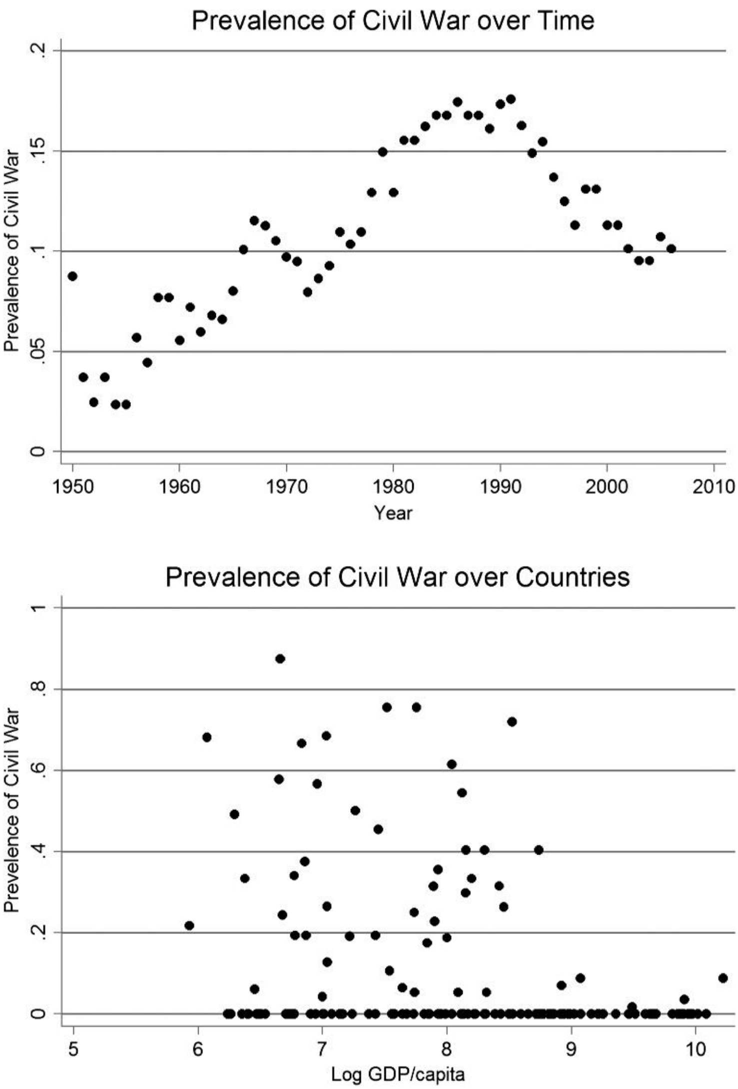

FIGURE I

(Continued)

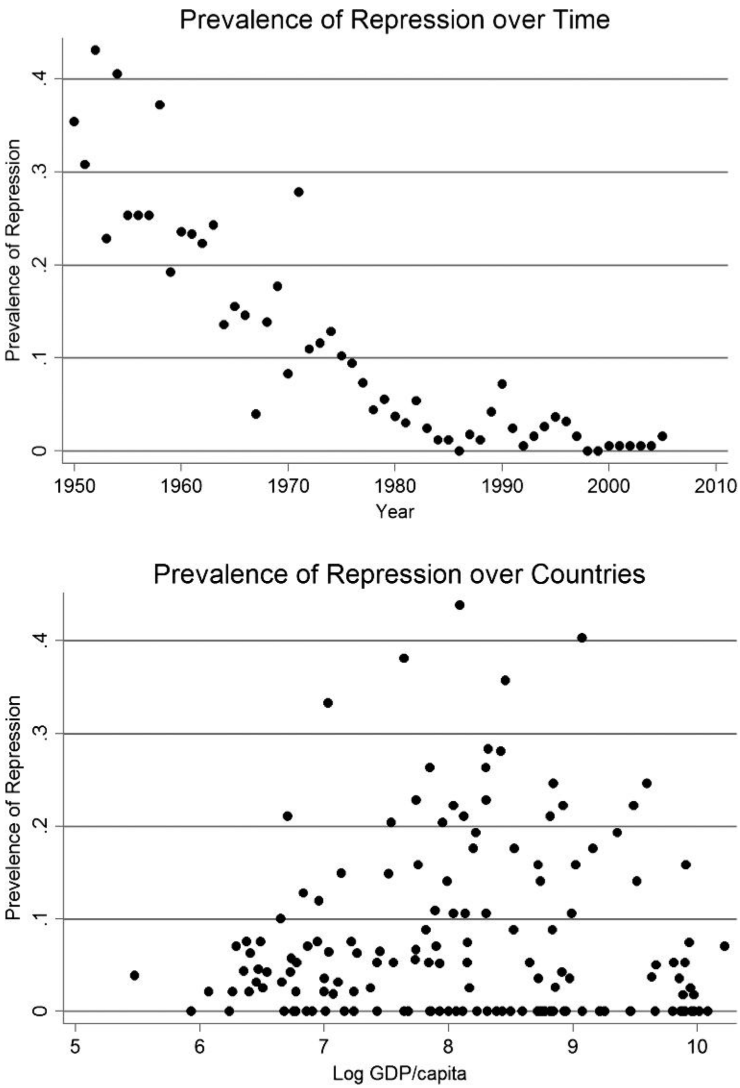

FIGURE I

Prevalence of Civil War and Repression

Figure I shows the worldwide trend in the development of purges. Interestingly, up to the early 1990s, this prevalence of repression series is almost a mirror image of the civil war series in the graph. When we plot the prevalence of repression by country against the level of GDP in 1980, it is striking that repression is most common in countries with higher income than in those where civil war is prevalent. Of course, outright conflicts and government repression come in different forms. Here, we focus on large-scale and serious manifestations of violence: civil war rather than civil conflict, and major rather than minor acts of government repression.[1]

The main contributions of the article are threefold. First, we develop the theoretical link between civil war and repression and show that they have common roots, especially as the product of noncohesive political institutions.[2] The theoretical framework allows us to study jointly the determinants of one-sided and two-sided violence. Second, we show how the theory can be used as a guide for measurement and formulating an empirical strategy. Third, we present econometric estimates which shed light on our theoretical predictions.

Our analysis builds on earlier research, which has developed both in its scope and its sophistication. By now, there exists a large amount of work by political scientists and economists on the causes of civil war. This literature has progressed from mainly cross-sectional inference using country-level data to panel-data studies, which exploit within-country variation—see the survey by Blattman and Miguel (2009). A largely independent literature, surveyed in Davenport (2007), has explored the determinants of government repression and violations of human rights. The main focus in both these strands of work has been on exploring empirical regularities, searching in some cases for credibly exogenous sources of variation. Links between theoretical models of conflict and violence are limited—both Blattman and Miguel (2009) and Davenport (2007) lament the fact that so few empirical findings forge links between the theory and data.[3]

The article is organized as follows. Section II develops our model where an incumbent government and an opposition group each can make an investment in political violence. The resulting conflict game is embedded in a public policy setup, where the

1. We also ignore other forms of violence such as riots and political intimidation. See, for example, Urdal (2008) or Bohlken and Sergenti (2011) for some recent work on how such violence relates to economic factors in India.

2. In a short previous paper, Besley and Persson (2009a), we brought out some of these ideas in a simple linear example.

3. There are certainly exceptions, however, such as Dube and Vargas (2008), who build explicitly on the theoretical framework developed by Dal Bó and Dal Bó (2011). See also Fearon (2008).

ruling group in each period controls the government budget, which can be used either for public goods or for redistribution between the two groups. This framework is capable of generating peace, repression (one-sided violence), and civil war (two-sided violence) as alternative equilibrium outcomes. We identify specific conditions on the conflict technology, under which these three conflict states are *ordered* in a latent variable, which summarizes the main determinants of conflicts: the level of resource rents, aid or other forms of income to the state, the level of wages, and the level of public goods provision. Importantly, however, our theory predicts an influence of these determinants on violence *only if* political institutions are noncohesive. Our theoretical results are summarized in two propositions and four corollaries.

Section III discusses how the theoretical predictions can guide empirical testing under specific assumptions about which elements of the theory are observable to the econometrician. This provides a particular take on the pitfalls in using cross-sectional variation in the data as the main source of variation. Following other recent contributions that have exploited panel data, we argue that a more credible way to identify causal links is to rely on within-country variation in the drivers of conflict—in our case, as suggested by the theory. Moreover, the theoretical framework naturally gives way to an ordered logit specification, with fixed country and year effects, for the states of peace, repression, and civil war.

Section IV describes our data on political violence, shocks, and political institutions, and presents our empirical results. We exploit two sources of, arguably, exogenous variation in the determinants identified by our theory: natural disasters—for negative shocks to wages and positive shocks to aid flows—and membership in the UN Security Council during the Cold War—for positive shocks to aid flows. Our empirical estimates are entirely consistent with the specific model predictions. Thus, natural disasters and Cold War Security Council membership both raise the probability of political violence—that is, civil war or repression—but only under noncohesive political institutions. In line with our theoretical priors, it is the combination of shocks and weak institutions that drive the empirical findings. In terms of specific mechanisms, it appears that most of the variation in political violence is tied to variation in aid flows.

Overall, the article begins to integrate several separate literatures, with theoretical as well as empirical work. Although we do not provide a general literature review in one place, we relate our approach to the existing literature as we go along. An appendix includes the proof of our main theoretical result. Section V concludes.

II. Theoretical Framework

Our aim is to build a simple and tractable model that serves as a useful guide to how measurable economic and political factors determine the probability of observing political violence.

Models that generate outright conflict as an equilibrium outcome rely on either imperfect information or inability of the parties to commit. The key friction in our model is of the second type: an inability of any prospective government to offer postconflict transfers credibly, and an inability of potential insurgents to commit not to using their capacity to engage in conflict.

There are two groups: A and B, each of which makes up half the population. Time is infinite and denoted by $t = 1, \ldots,$ although we drop the time index in this theoretical section. One generation is alive at each date and is labeled by the date at which it lives. The model has no state variables, so the dynamics come only from three stochastic variables—wages, the value of public goods and of exogenous income (from natural resources or aid)—which are identically and independently distributed over time.

At the beginning of each period, members of the group that held power at the end of the previous period inherit a hold on the incumbent government, denoted by $I \in \{A, B\}$. The other group makes up the opposition, denoted by $O \in \{A, B\}$. Power can be transferred by peaceful means. But to raise its probability to stay in power, the incumbent group can invest in violence, an investment denoted by L^I—think about mounting an army. To try to take over the government, the opposition can also invest in violence with armed forces L^O—think about mounting an insurgency. The conflict technology is discussed later. Whether power is transferred peacefully or through armed conflict, the winner becomes the new incumbent and the loser the new opposition, denoted by $I' \in \{A, B\}$ and $O' \in \{A, B\}$.

The new incumbent gets access to existing government revenue, from for example, aid, natural resources, or taxes, which is denoted by R. The exogenous revenue stream is divided between spending on general public goods G and transfers to the incumbent $r^{I'}$ and the opposition $r^{O'}$. Revenues are stochastic and drawn afresh each period from $R \in [R_L, R_H]$. The precise timing of these different events/decisions are spelled out shortly.

Individual incomes and utility. Individuals supply labor in a common labor market to earn an exogenous wage w. Like revenues, wages are stochastic and distributed in the interval: $w \in [w_L, w_H]$. Individuals have quasi-linear utility functions:

$$V^J = \alpha H(G) + c^J, \tag{1}$$

where c^J is private consumption by group $J \in \{I', O'\}$ and G is the level of public goods provided, with the parameter α reflecting the value of public goods. The function $H(\cdot)$ is increasing and concave. The value of public goods is stochastic with $\alpha \in [\alpha_L, \alpha_H]$.

The government budget constraint in any period can be written

$$R - \sum_{J \in \{I', O'\}} \frac{r^J}{2} - G - wL^I \geq 0, \tag{2}$$

where L^I denotes the size of the incumbent's army, which is thus financed out of the public purse.

Institutions. As mentioned, power can be transferred between groups peacefully, or as a result of groups making investments in violence, L^J. The probability that group O wins power and becomes the new incumbent I' is

$$\gamma\left(L^O, L^I; \xi\right), \tag{3}$$

which depends on the resources devoted to fighting and a parameter vector, ξ. We use γ_O and γ_I to denote first derivatives with regard to the first and second arguments of the function in Equation 3, respectively, with second derivatives γ_{II}, γ_{OO}, and γ_{IO} defined correspondingly.

The function 3 is a *contest function* of the kind used in the existing theoretical literature on conflict (see, for example, Dixit 1987 and Skaperdas 1996 which surveys the use of contest functions and their properties). We assume that the function $\gamma(\cdot)$ is increasing in its first argument and decreasing in the second. In this notation, $\gamma(0, 0; \xi)$ is the probability of a peaceful transition of power between the groups. Later, we make further specific assumptions on the properties of the function in Equation 3.

Each group (when in opposition) has the power to tax/conscript its own citizens to finance a private militia to mount an insurgency. We denote this capacity by X (so $L^O \leq X$), which is common to the two groups, so that neither has a greater intrinsic capability to fight. This unified-actor formulation sweeps aside the interesting issue of how it is that an opposition can solve the collective action problem in organizing violence.

Political institutions are assumed to constrain the possibilities for incumbents to make transfers to their own group. To capture this as simply as possible, assume that an incumbent government must give $\sigma \in [0, 1]$ to the the opposition group, when it makes a transfer of 1 to its own group, implying that $r^{O'} = \sigma r^{I'}$. Given this assumption, we use the government budget constraint (assuming it holds with equality) to obtain:

$$r^{I'} = 2\left(1-\theta\right)\left[R - G - wL^I\right], \tag{4}$$

where $\theta = \frac{\sigma}{1+\sigma} \in \left[0, \frac{1}{2}\right]$. Throughout, we interpret a higher value of the opposition's share of transfers, θ, as reflecting more representative, or cohesive, political institutions. The real-world counterparts of a high θ may be more minority protection through a system of constitutional checks and balances, through a parliamentary form of government, or through a proportional electoral system. If $\theta = \frac{1}{2}$, then transfers are shared equally across the two groups. Thus, we can think of θ as an institutionalized ability to make commitments not to expropriate the opposition; θ closer to (further from) one half represents a case of stronger (weaker) political institutions.

Timing. The following timing applies to each generation:

1. The value of public goods α, the wage rate w, and revenues (natural resource rents or aid) R are realized.
2. Group I and group O simultaneously choose the sizes of their armies.
3. Group I remains in office with probability $1 - \gamma\left(L^O, L^I; \xi\right)$.
4. The winning group becomes the new incumbent I' and determines policies, that is, spending on transfers $\left\{r^J\right\}_{J\in\{I',O'\}}$ and public goods G.
5. Payoffs are realized, consumption takes place, and the currently living generation dies.

We next solve the model by working backward to derive a subgame perfect equilibrium.

Equilibrium policies. Suppose we have a new incumbent at stage four. Then, using Equation 4, the optimal level of public goods is determined as:

$$G = \arg\max_{G \geq 0} \left\{ \alpha H(G) + 2(1-\theta)\left[R - G - wL^I\right] + w \right\}. \tag{5}$$

Defining $\widehat{G}(z)$ by

$$H_G\left(\widehat{G}(z)\right) = \frac{1}{z},$$

we can record the policy solution as:

LEMMA 1. For given (α, w, R), public goods are provided as:

$$G = \min\left\{ \widehat{G}\left(\frac{\alpha}{2(1-\theta)}\right),\ R - wL^I \right\}.$$

There are two cases. If α is large enough and/or R small enough, all public spending goes on public goods with any incremental revenues also spent on public goods. Otherwise, the optimal level of public goods is interior and increasing in α and θ. Intuitively, transfers to the incumbent's own group become more expensive as θ increases. Since $\theta \leq \frac{1}{2}$, public-goods provision is below $\widehat{G}(\alpha)$, the choice of a utilitarian planner. With an interior solution for G, any residual revenue is spent on transfers, which are distributed according to the θ-sharing rule.

Political violence. We now study the possibility of conflict at stage two, looking for an equilibrium in which the opposition decides whether to mount an insurgency and the incumbent government chooses how to respond. As we will show, the equilibrium has three possible regimes. In the first, no resources are invested in violence by either side, that is, peace prevails. In the second, there is no insurgency, but the government uses armed forces to repress the opposition and thereby increase its chances of remaining in power. In the third case, there is outright conflict, where both sides are investing in violence and committing military resources to a civil war.

Using the results in the last subsection, it is easy to check that the expected payoff of the incumbent is:

$$\begin{aligned} \hat{V}^I\left(\alpha, w, R; L^O, L^I\right) &= \alpha H(G) + w + [(1-\theta) - \gamma\left(L^O, L^I; \xi\right) \\ &\quad \times (1-2\theta)]2\left[R - G - wL^I\right]. \end{aligned} \tag{6}$$

The key term is $[(1-\theta)-\gamma\left(L^{O},L^{I};\xi\right)(1-2\theta)]$, the weight the incumbent attaches to end-of-period transfers. This is the average share of the incumbent, $(1-\theta)$, given the institutional restriction on transfers, minus the probability that the opposition takes over times the "extra" share, $(1-2\theta)$, that an incumbent captures of the redistributive pie.

For the opposition group, we have

$$(7)\qquad \hat{V}^{O}\left(\alpha,w,R;L^{O},L^{I}\right)=\alpha H\left(G\right)+w\left(1-L^{O}\right) + [\theta+\gamma\left(L^{O},L^{I};\xi\right)(1-2\theta)]2\left[R-G-wL^{I}\right],$$

where $[\theta+\gamma\left(L^{O},L^{I};\xi\right)(1-2\theta)]$ is the opposition's expectational weight on transfers.

These payoff functions expose the asymmetry between the incumbent and opposition in terms of financing the army. The incumbent's army is publicly financed and increasing the size of it reduces future transfers. The opposition must finance any insurgency out of the group's own private labor endowment given the power to tax its own citizens.

The two payoff functions also express the basic trade-off facing the two parties. On one hand, higher armed forces have an opportunity cost. On the other hand, for given armed forces of the other party, they raise the probability of capturing or maintaining power to take advantage of the monopoly on allocating government revenue.

To solve for the equilibrium level of conflict, define $Z=\frac{R-G}{w}$, a stochastic variable that depends on realizations of the vector (α,w,R). This is the level of adjusted and uncommitted government revenues, specifically the ratio of the maximal redistributive pie (what can be spent on transfers, given equilibrium public goods provision) to the real wage. The equilibrium can then be described by two threshold values for Z, the size of the wage-adjusted redistributive pie, above which the incumbent and opposition find it worthwhile to expend positive resources on fighting.

We characterize a Nash equilibrium $\left(\hat{L}^{I},\hat{L}^{O}\right)$ of the conflict game in pure strategies, where

$$\hat{L}^{I}=\arg\max 2w\left\{\left[1-\theta-\gamma\left(L^{O},L^{I};\xi\right)(1-2\theta)\right]\left[Z-L^{I}\right]\right\}$$

for the incumbent and

$$\hat{L}^O = \arg\max w \left\{ 2\left[\theta + \gamma\left(L^O, L^I; \xi\right)(1-2\theta)\right]\left[Z - L^I\right] - L^O \right\}$$

for the opposition.

We first state a simple result:

PROPOSITION 1. As $\theta \to \frac{1}{2}$, there is always peace.

Proof. When $\theta \to \frac{1}{2}$, the expressions for $\hat{L}^I$ and $\hat{L}^O$ are decreasing in L^I and L^O, respectively. ■

Intuitively, when θ is close to one half, there is no gain from fighting since institutions constrain the use of the state to give both groups basically the same share of any transfers regardless of who is in office. Thus, there is no point in expending costly resources to struggle for power. This gives a simple account for why we predominantly observe political violence in countries with noncohesive political institutions.

To study the Nash equilibrium when institutions do not make a country conflict proof, we make the following assumption on the conflict technology:

ASSUMPTION 1. For all $L^O \in [0, X]$ and $L^I \in [0, Z]$, the conflict technology satisfies:

a. $\gamma \in (0,1)$, $\gamma_O > 0, \gamma_I < 0$, $\gamma_{OO} < 0$, $\gamma_{II} > 0$,

b. $\dfrac{-\gamma_I(0,0;\xi)}{\gamma_O(0,0;\xi)} \geq 2\left[1 - \gamma(0,0;\xi)\right]$, and

c. $\dfrac{\gamma_I \gamma_{OO}}{\gamma_O} \geq \gamma_{IO} \geq \dfrac{\gamma_O \gamma_{II}}{\gamma_I}$.

Condition a just says that neither group can ever be fully certain of holding power, and that fighting always has positive returns for both groups, albeit at a decreasing rate. Property b ensures that the incumbent has a higher marginal return to fighting, when both parties do not invest in violence, and/or the incumbent faces a sufficiently high probability of losing power peacefully. Finally, c restricts the extent of any strategic complementarities or substitutabilities in the conflict technology.

Using Assumption 1, we have the following characterization of the Nash equilibrium.

PROPOSITION 2. If Assumption 1 holds and $\theta < \frac{1}{2}$, there exist two thresholds $Z^I(\theta;\xi)$ and $Z^O(\theta;\xi)$ with $Z^I_\theta, Z^O_\theta > 0$ and

$$Z^I(\theta;\xi) = \frac{\frac{(1-\theta)}{(1-2\theta)} - \gamma(0,0;\xi)}{-\gamma_I(0,0;\xi)} < Z^O(\theta;\xi),$$

such that:

1. For $Z \leq Z^I$ there is peace with $\hat{L}^O = \hat{L}^I = 0$,
2. For $Z \in \left(Z^I, Z^O\right)$, there is repression with $\hat{L}^I > \hat{L}^O = 0$,
3. For $Z \geq Z^O$ there is civil conflict with $\hat{L}^I, \hat{L}^O > 0$.

Moreover, the level of violence, whenever positive, is increasing in Z for both the incumbent and the opposition groups.

Proof. See the Appendix. ■

The proposition describes three cases. When Z is below Z^I, no conflict erupts as both the incumbent and the opposition accept the (probabilistic) peaceful allocation of power, where the opposition takes over with probability $\gamma(0,0;\xi)$. When $Z \in \left[Z^I, Z^O\right]$, the government invests in violence to increase its survival probability, but the opposition does not invest in conflict. Finally, when $Z > Z^O$, the opposition mounts an insurgency, which is met with force by the incumbent group.

Discussion. Though the result in Proposition 2 is intuitive, it is important to assess the specific assumptions used in deriving it. Assumption 1b rules out an undefended insurgency. It says that the return to fighting is strong enough for the incumbent, given the threat of political transition under peace. If this assumption does not hold, we may have a range of Z where the incumbent does not bother to fight the opposition when it rebels. This might be true, for instance, if $\gamma(0,0;\xi)$ is very close to 0 and $\frac{-\gamma_I(0,0;\xi)}{\gamma_O(0,0;\xi)}$ is close to 0 so that the incumbent is not very threatened by a transition and/or has low competence in defending against it. We find it natural to rule out undefended insurgencies, because we think such phenomena are rare. But they could be encompassed as a theoretical possibility in our framework.

Assumption 1c guarantees that the fighting propensities of both incumbent and opposition increase in the size of the prize, measured by Z. Given that a civil war has started, this ensures that increasing Z does not make either party give up. This will be true as long as the marginal return to fighting is not strongly affected by the fighting decisions of the other group, placing bounds on γ_{IO}, not allowing a positive or negative cross-partial that is too large.[4]

While we have kept the contest function general, the model works with a number of reasonable and widely used specific contest functions. For example, it works with the popular ratio formulation (see Tullock 1980 and Skaperdas 1992) if

$$\gamma\left(L^O, L^I; \xi\right) = \frac{\xi L^O}{\xi L^O + L^I} \ ,$$

where parameter $\xi \geq 1$.[5] Similarly, we can use the logistic formulation (see Hirshleifer 1989) if

$$\gamma\left(L^O, L^I; \xi\right) = \frac{\exp[\xi_O L^O]}{\exp[\xi_O L^O] + \exp[\xi_I L^I]} \ ,$$

and $\xi_I \geq \xi_O$, or the semi-linear formulation:

$$\gamma\left(L^O, L^I; \xi\right) = \gamma_O + \xi_1\left[h\left(L^O\right) - \xi_2 h\left(L^I\right)\right] \ ,$$

where $h(\cdot)$ is an increasing concave function, with $h(0) = 0$, $h_L(0) > 0$, and $\bar{h} = \lim_{z\to\infty} h(z)$, capturing how investments in arms translate into violence, with parameter restrictions $\xi_1 > 0$, $\xi_2 \geq 1$ and $1 - \xi_1\bar{h} \geq \gamma_O \geq \max\left\{\frac{1}{2}, \xi_1\xi_2\bar{h}\right\}$.

Implications. Our results have some striking empirical implications when the logic of political violence is expressed as a function of latent variable Z. More precisely, our theory predicts an *ordering* in Z of the three states peace, repression, and civil war. This ordering is particularly interesting against the backdrop of Figure I, which suggests that repression and civil war have been substitutes, at least for some of the time and some of the countries, in the post-war period.

4. We could make the weaker assumption that $\frac{\partial}{\partial\lambda}\left(\frac{-\gamma_I(\lambda x,(1-\lambda)x)}{\gamma_O(\lambda x,(1-\lambda)x)}\right) \geq 0$ for $\lambda \in [0,1]$ and $x \geq 0$ which is implied by Assumption 1c. This amounts to saying that the conflict technology is quasi-concave, that is, has level sets that are convex in $\left(L^O, L^I\right)$ space. This makes total spending on conflict by the two parties monotonic in Z, but not necessarily the spending by each group. In economic terms, this could lead to a resumption of repression or undefended insurgency at high levels of Z as one group drops out of the fight.

5. By l'Hopital's rule:

$$\gamma(0,0;\xi) = \frac{\xi}{\xi+1}.$$

The Z variable summarizes several important determinants of violence, which we now bring out in a set of corollaries. We state these in terms of likelihoods, implicitly assuming that some factors are not only uncertain but also unobserved by an outside analyst. A more precise formulation of the empirical predictions—along precisely these lines—is found in Section III.

COROLLARY 1. Higher wages, w, reduce the likelihood that an economy will experience political violence, that is, be in repression or civil war, unless political institutions are cohesive (θ close to $\frac{1}{2}$).

The result follows from Proposition 2 by observing that w is the denominator of Z. Given the distributions of α and R, when w is higher the whole distribution of Z thus shifts to the left. Based on this, we can definitely say that higher wages make peace more likely (political violence less likely). We can also definitely say that civil war becomes less likely. But whether repression is more or less likely depends on relative densities (in the p.d.f. of Z). The qualifier at the the end of the corollary follows directly from Proposition 1.

Of course, this result reflects a higher opportunity cost of fighting at higher wages, and hence a lower net gain from winning a conflict to both parties. In the literature on civil war, this effect is well known at least since Grossman (1991) and has been emphasized, in particular, by Collier and Hoeffler (2004).[6] Here, we see that the result extends to political violence more generally. In the empirical literature, this opportunity cost channel is most often proxied by the level of income per capita. However, whether changes in income per capita are a good proxy for wage changes depends on the underlying source of the shock.[7]

COROLLARY 2. Higher natural resource rents, or other exogenous forms of income such as aid, a higher R, increase the likelihood that an economy will be in repression or civil war, unless political institutions are consensual (θ close to $\frac{1}{2}$).

6. Chassang and Padró i Miquel (2009) also describes a mechanism to model the impact of economic shocks on conflict.

7. In the two-sector conflict model of Dal Bó and Dal Bó (2011), for example, world price shocks drive real wages and returns to capital in opposite directions, producing an unclear correlation between wages and income per capita.

The corollary follows from Propositions 1–2, once we note that Z depends directly on the level of natural resource rents or exogenous income to government from any other source, like aid. The effect of resource rents has been emphasized in the empirical literature on civil war (see, e.g., Humphreys 2005 and the surveys in Ross 2004 and Blattman and Miguel 2009), but few papers have derived the theoretical result (one of the first is Aslaksen and Torvik 2006). As far as we know, the rent-seeking channel does not figure much in the literature on repression and human rights infringements.

COROLLARY 3. Higher spending on common-interest public goods, induced by higher α, reduces the likelihood that an economy will be in repression or civil war, unless political institutions are cohesive (θ close to $\frac{1}{2}$).

This follows because an increase in α raises G and hence reduces Z. To the best of our knowledge, this specific prediction of our model is new to theoretical models of civil war, since conflict models are typically not embedded in an explicit public finance context. At a general level, however, the broad selectorate framework in Bueno de Mesquita et al. (2003) considers the split of government revenue into public goods versus redistribution, as well as government repression and civil war, as endogenous outcomes. In their analysis, some institutional variation—such as a larger winning coalition within the selectorate—might produce a correlation between public goods and violence similar to the one entailed in Corollary 3.

While these three implications of the model all reflect variations in Z, other parameters will affect conflict by changing the two trigger points Z^O and Z^I. Such will be the case with parameters of the conflict technology ξ, but to sort these out requires additional specific assumptions. However, we directly obtain a result concerning the effect of political institutions.

COROLLARY 4. Political institutions with more checks and balances and more minority representation, a higher value of θ, decrease the likelihood of observing repression or civil war (in the range of θ for which the equilibrium is not necessarily peaceful).

This follows by observing that $Z^O(\theta;\xi)$ and $Z^I(\theta;\xi)$ are both increasing functions of θ. Intuitively, more cohesive institutions make control of the state less valuable, and thus shift up the point at which Z triggers violence both for the incumbent and the opposition. Many of the papers in the civil war and repression literatures discuss and attempt to estimate the dependence of violence on political institutions, but typically as a direct affect. However, Propositions 1–2 also have the joint implication that Corollaries 1–3 should only hold in societies and times where θ—the minority protection or representation embedded in political institutions—is below a certain lower bound. As far as we know, this specific theoretical insight from our model is also new.

III. From Theory to Econometric Testing

In this section, we discuss how our theory can inform the empirical study of political violence. Although our model is extremely simple, it does give a transparent set of predictions for how parameters of the economy and the polity shape the incidence of violence. A clear advantage of beginning from a well-defined theory is that we may clarify and evaluate the assumptions made en route to empirical testing. Specifically, we must take a stance on which variables and parameters are measurable in the data—that is, which are observable and which are not—as well as which variables and parameters to treat as fixed (at the country level) rather than time varying.

Measurement, observability, and likelihoods. Our data are in panel form for countries and years from 1950 onward. Hence, consider country c at date t. Shortly we discuss how we can use readily available sources of data to decide whether that country-year is characterized by peace, repression, or civil war. When it comes to the components of the latent index variable $Z_{c,t}$, we will argue that for each country, we can find time-varying correlates of $w_{c,t}$ and $R_{c,t}$ which we also discuss shortly.

However, we cannot measure variations in public goods, as induced by time-varying parameter $\alpha_{c,t}$, because we are unable to gather data on public goods provision for a large enough sample of countries during a long enough time. Because of this, we will not be able to test Corollary 3. Given the model, let $\varepsilon_{c,t} = \widehat{G}(\frac{\alpha_{c,t}}{2(1-\theta_c)}) - G_c$ be the country-specific randomness in public goods provision, where G_c is the country-specific unobserved mean of G. Then, $\varepsilon_{c,t}$ will have some country-specific c.d.f. $F^c(\varepsilon)$ on finite support $\left[\widehat{G}\left(\frac{\alpha_L}{2(1-\theta_c)}\right) - G_c, \widehat{G}\left(\frac{\alpha_H}{2(1-\theta_c)}\right) - G_c\right]$. As for the other parameters of the model, we will treat them as constant over time. Finally, although we will be able to observe proxies for the cohesiveness of political institutions, θ_c, we do not readily observe parameters of the conflict technology, ξ_c.

Using Proposition 2 and the definition of Z, we can then express the condition for civil war in country c at date t as

$$Z_{c,t} - Z^O\left(\theta_c;\xi_c\right) = \frac{R_{c,t}}{w_{c,t}} - Z^O\left(\theta_c;\xi_c\right) - \frac{G_c}{w_{c,t}} - \frac{\varepsilon_{c,t}}{w_{c,t}} \geq 0.$$

Under our assumptions, the conditional probability for an outside researcher to observe conflict in country c at date t is thus given by:

$$F^c(R_{c,t} - Z^O\left(\theta_c;\xi_c\right) w_{c,t} - G_c)\,. \tag{8}$$

As predicted by the theory, a higher value of $R_{c,t}$ or a lower value of $w_{c,t}$ both raise the likelihood of observing civil war, *provided* that θ is not close to $\frac{1}{2}$.[8]

By similar reasoning, the likelihood of observing peace is

$$1 - F^c(R_{c,t} - Z^I\left(\theta_c; \mathrm{s}\,\xi_c\right) w_{c,t} - G_c)\,, \tag{9}$$

while the likelihood of observing repression is

$$F^c(R_{c,t} - Z^I\left(\theta_c;\xi_c\right) w_{c,t} - G_c) - F^c(R_{c,t} - Z^O\left(\theta_c;\xi_c\right) w_{c,t} - G_c)\,. \tag{10}$$

As explained in Section II, the theory gives us distinct predictions how changes in $R_{c,t}$ and $w_{c,t}$ shift the distribution of index variable $Z_{c,t}$ and thereby the likelihood of observing peace, while the predictions regarding the conditional probability of observing recession hinge on the relative densities of F^c. In other words, we have specific predictions about two margins: that between civil war and non–civil war (peace *cum* repression), and that between peace and political violence (repression *cum* civil war).

Another informative way of interpreting expressions 8–10 is that they define the relative probabilities of the three ordered states of violence. This strongly suggests that the most straightforward way of confronting the theory with data would be to estimate a fixed-effect ordered logit driven by variables that shift

8. Formally, as θ approaches $\frac{1}{2}$, Z^I and hence $Z^O > Z^I$ approach infinity. Given the finite support for the distributions of α, w, and R, the maximum of F^c, namely $F^c(R_H - Z^O w_L - G_c)$ is thus equal to 0.

the country-specific distribution of $Z_{c,t}$ given the country-specific thresholds $Z^I(\theta_c;\xi_c)$ and $Z^O(\theta_c;\xi_c)$.

Cross-country versus within-country variation. What kind of variation in the data should we use to test the model predictions? A good deal of the empirical civil war literature, and virtually all of the empirical repression literature, estimates the probability of observing violence from cross-sectional data sets. Expressions 8–10 illustrate clearly why this may not be such a good idea. Cross-sectional data replace the time-varying variables $R_{c,t}$ and $w_{c,t}$ with their cross-sectional means R_c and w_c. But this makes statistical inference a hazardous exercise, because it runs the risk of confounding the cross-country variation in these variables with cross-country variation in the unobserved parameters G_c and ξ_c, something that could seriously bias and invalidate the estimates.

It is more rewarding to exploit within-country variation in panel data, as in the cross-country panel studies of civil war in Africa by Miguel, Satyanath, and Sergenti (2004) or Bruckner and Ciccone (2008), and the within-country panel studies of civil war by Deininger (2003) for Uganda or Dube and Vargas (2008) for Colombia. For instance, estimating a specification for the likelihood of observing civil war, with fixed country effects, is equivalent to evaluating

$$F^c(R_{c,t} - Z^O(\theta_c;\mathrm{s}\,\xi_c)\,w_{c,t} - G_c) - E\{F^c(R_{c,t} - Z^O(\theta_c;\xi_c)\,w_{c,t} - G_c)\}, \tag{11}$$

that is, the difference between the conditional and the unconditional probability of civil war. Proceeding in this way identifies the effect of resource rents/aid flows $R_{c,t}$ and wages $w_{c,t}$ on the incidence of civil war exclusively from the within-country variation of these variables. Any impact of their average values and time-invariant parameters in each country are absorbed by the country fixed effect.

Given the important and irregular time trends in the prevalence of civil war and repression in Figure I, it is also essential to allow for global shocks, which hit all countries in a common way, through year fixed effects (time indicator variables). The trends in violence are then picked up in a flexible (nonparametric) fashion, and we only use the country-specific yearly variation relative to world year averages for identification.

Our specification should also take into account that the predictions about shocks are conditional on the value of θ_c. Let $\Theta_c = 1$

if political institutions have strong checks and balances (i.e., θ_c close to $\frac{1}{2}$) in country c in the period of our data, and equal to 0 otherwise. We then model the index function in 11 as:

$$(12) \qquad R_{c,t} - Z^O\left(\theta_c; \xi_c\right) w_{c,t} - G_c = a_c\left(\Theta_c\right) + a_t\left(\Theta_c\right) + b\left(\Theta_c\right) \widetilde{Z}_{c,t} \ ,$$

where $a_c\left(\Theta_c\right)$ is a country fixed effect, $a_t\left(\Theta_c\right)$ are year dummies, and $\widetilde{Z}_{c,t}$ are time-varying regressors that reflect changes in $R_{c,t}$ and $w_{c,t}$. The theory predicts that the parameter of interest, $b\left(\Theta_c\right)$, is heterogeneous with respect to Θ_c, in particular, that $b(0) > b(1) = 0$. To test this prediction, we estimate a model that allows for separate slope coefficients for weakly and strongly institutionalized countries.[9]

IV. Data and Results

In this section, we first describe our data and then present our empirical results.

Data: political violence and political institutions. A large body of literature looks at the determinants of civil war.[10] In this article we mainly use the ACD civil war incidence measure, starting in 1950.[11] It takes a value of 1 if—in a given country and year—the government and a domestic adversary are involved in a conflict, which claims a cumulated death toll of more than 1,000 people. As mentioned in the introduction, over 10% of all country-years in the 1950–2005 period are classified as civil war in our sample.[12] Because we want to focus of large-scale political violence, we do not exploit the alternative oft-used incidence of civil conflict (also from the ACD), which only requires a cumulated death toll of 25 people.

9. In the specifications reported in Tables I–II we impose $a_c(1) = a_c(0)$ and $a_t(1) = a_t(0)$. However, the results hold up when we allow for separate country and time effects by estimating the model on separate subsamples, that is, with $\Theta_c = 1$ and $\Theta_c = 0$.

10. There are a number of issues involved in the coding of conflicts into civil wars. See Sambanis (2004) for a thorough discussion about different definitions that appear in the empirical literature.

11. Specifically, we use the variable "Incidence of intrastate war" in the UCDP/PRIO Armed Conflict Dataset v.4-2007, covering the years 1946–2006.

12. An alternative measure is available in the Correlates of War (COW) database, but this only runs up to 1997. Given that one of our independent variables relies on Cold War and post-Cold War experience, the COW variable would only allow for 8, as opposed to 16 observations, in the post-Cold War era.

To measure repression, we use a measure from Banks (2005) that counts up purges: systematic murders and eliminations of political opponents by incumbent regimes. We create an indicator that is equal to 1 in any year when purges exceed 0. In the 1950–2005 period, on average 7% of country-years are classified as being in a state of repression, but not in civil war.[13]

Based on these two measures, we construct our ordered variable of political violence. Specifically—and without loss of generality, as only the ordinal ranking matters—we assign a value of 0 to peace, a value of 1 to repression in the absence of civil war, and a value of 2 to civil war.[14]

Are these three states naturally ordered in the data, as in the theory? For income per capita, the answer is a clear-cut yes. Peaceful country-years have an average GDP per capita of $4,365, repressing countries are poorer with $2,503 per capita, and those in civil war are the poorest with average incomes of $1,789.

We construct two indicator variables to capture cohesive political institutions, corresponding to Θ_c in Section III. Our core measure is based on the assessment of executive constraints in the Polity IV data set.[15] We believe this variable best captures the thrust of θ in our theory. Executive constraints are coded annually from 1800 or from the year of independence. We do not exploit the high-frequency time variation in this variable, however, as we are concerned that changes are likely to be correlated with the incidence of political violence.[16] This means that we leave a test of Corollary 4 for future work.

13. An alternative would be to exploit the commonly used Political Terror Scale based on the reports on human rights violations by the U.S. State Department and Amnesty International. This variable is only available from 1976, however, which cuts short the Cold War period that we can exploit. Moreover, as shown by Qian and Yanagizawa (2009), Security Council membership during the Cold War period may have affected the way the U.S. State Department reported on human rights in allied and nonallied countries.

14. To be precise, we begin from two underlying variables: civil wars as coded in the ACD and the purges variable in Banks (2005). We construct a binary variable based on the latter depending on whether there are some purges in a country at a given date. Since 1950, we have 4841 country-year observations with neither civil war nor government purges. There are 90 observations where there is both a civil war and some purges, 714 observations where there are civil wars but no purges, and 425 observations where there are purges but no civil war. This yields 1,229 observations with some violence and 804 with civil war.

15. In the Polity IV this is variable “XCONST”.

16. Besley and Persson (2011) formulates a model where political violence and political institutions are both endogenous.

To construct a time-independent measure of Θ_c, we adopt a somewhat conservative approach. First, we evaluate the presample evidence, measuring the fraction of years for which a country had the highest score (of seven) for executive constraints before 1950. Then, we compute the fraction of years for which a country has the top score over the sample period. A country is deemed to have strong political institutions, $\Theta_c = 1$, if the fraction in the presample period is above 0, *and* the fraction in the sample period is greater than 0.6. This definition classifies about 18% of countries into cohesive institutions.[17] Marginal changes in the classification criteria have little effect on the results.

Using this variable, we uncover a striking regularity across political regimes. For countries with cohesive institutions, 93% of the annual observations are peaceful with 3.7% in repression and 2.8% in civil war. For countries with noncohesive institutions, these figures are 77%, 8%, and 15%, respectively. Such a difference between the two groups in the unconditional probability of observing political violence is in line with our theory.

As a robustness check, we use an alternative classification of political institutions based on the prevalence of parliamentary democracy. While high executive constraints are associated with stiffer checks and balances on the government, the alternative measure is intended to capture greater representativeness.[18] We define it analogously, namely as the result of having had a positive prevalence of parliamentary democracy before 1950, and a minimum prevalence of 0.6 between 1950 and 2005.

Data: $\widetilde{Z}$-shocks. To test the specific model predictions with the specification in Equation 12, we still need credibly exogenous variation in the time-varying regressors $\widetilde{Z}_{c,t}$. We use two variables for this purpose.[19] The first is a measure of natural disasters,

17. The 26 countries are Australia, Austria, Belgium, Canada, Costa Rica, Denmark, Estonia, Finland, Germany, Greece, Iceland, Ireland, Israel, Italy, Japan, Latvia, Lithuania, Luxembourg, Netherlands, New Zealand, Norway, Sweden, Switzerland, South Africa, United Kingdom, and the United States.

18. See Persson, Roland, and Tabellini (2000) or Aghion, Alesina, and Trebbi (2004) for theoretical arguments, and Persson and Tabellini (2003) for empirical evidence.

19. An earlier version of the article also relied on commodity-price variation in world markets, measured through a country-specific export-price index, to gauge exogenous variation in resource rents.

constructed from the EM-DAT data set.[20] Specifically, we define a variable that adds together the number of extreme temperature events, floods, slides and tidal waves in a given country and year.[21] Then, we then create a binary indicator variable, set equal to 1 if a country experiences any such event. We expect this binary variable to negatively affect the real wage $w_{c,t}$. Consistent with this, having at least one natural disaster is associated with a 2.5% reduction in income per capita in the same country year. But part of this could be a productivity effect working through destruction of capital.[22] Of course, a natural disaster is also likely to trigger international aid flows. In terms of our theory, this corresponds to a positive shock to $R_{c,t}$, which affects the likelihood of violence in the same direction as a negative shock to $w_{c,t}$.

As a second source of exogenous variation, we use the revolving memberships in the UN Security Council (for nonpermanent members). We expect membership to raise a country's geopolitical importance and therefore its susceptibility to receive international aid from important countries, corresponding to positive shocks to $R_{c,t}$. Indeed, Kuziemko and Werker (2006) find that U.S. aid flows depend on Security Council membership. Similar incentives are likely to have applied to other permanent Security Council members. Of course, Security Council memberships may also change a country's international accountability, reducing the likelihood that its government engages in violence. Therefore, we mainly exploit the interaction between membership and time, allowing for a different effect before and after the fall of the Berlin Wall. In particular, we expect the strategic aid motives to be considerably stronger in the period before 1990, because of the stronger geopolitical tensions during the Cold War.[23]

20. Following an early paper by Drury and Olson (1998), Nel and Righarts (2008) investigate the association between different forms of natural disasters and civil conflict.

21. Specifically, we added together the variables "flood," "etemp," "slides," and "wave." Some other EM-DAT coded disaster events, such as epidemics, are not used because they may be endogenous to civil wars.

22. Recalling the discussion after Corollary 1, we could think about a natural disaster as a negative TFP shock plus (stochastic) depreciation of part of the capital stock. This would cut wages and perhaps the return to capital. In a more elaborate model, a lower return to capital may also cut the opportunity cost for engaging in conflict and so have a similar effect on conflict propensity as a lower return to labor.

23. See Bates (2008) for a discussion of how the Cold War affected government in Africa. Possibly, Cold War Security Council membership may affect

To explore the importance of these channels, we use data on total international aid disbursements from Organisation for Economic Co-operation and Development (OECD) countries, and on GDP per capita from the Penn World Tables.[24]

Basic results. Table I includes our core results.

In column (1), we present estimates from a fixed-effect ordered logit, a specification that is suggested by the theory. We implement this method of estimation using an approach proposed by Ferrer-i-Carbonell and Frijters (2004).[25] In addition to the (country and year) fixed effects, the specification includes our three exogenous variables. The panel for the estimation includes the 97 countries that have experienced some kind of political violence since 1950 (for the others, the fixed effect perfectly predicts the absence of violence). Column (1) shows that all three variables of interest are statistically significant: having a natural disaster is positively correlated with political violence, while being a member of the Security Council is negatively correlated with violence, except during the Cold War when the correlation is positive. The effect of having a natural disaster is nontrivial in magnitude: the point estimate corresponds to a little more than four percentage points higher probability of observing violence, given a sample average of about 17%. The effect of Security Council membership is of similar magnitude, predicting a four percentage point lower probability of political violence.

conflict through a different channel, namely, the provision of military aid raising the government's capability to fight. In the simple semi-linear conflict model mentioned in Section II, a higher value of ξ_2 can readily be interpreted as the incumbent's advantage in fighting. One can show that Z^I (the incumbent's trigger point) is decreasing in ξ_2, while Z^O (the opposition's trigger point) is increasing in ξ_2. Adding this channel to the effect of a higher Z via regular aid would mean that Cold War Security Council membership definitely should raise the likelihood of political violence, whereas it might raise or cut the likelihood of civil war.

24. More precisely, for aid we use the variable "Official Development Assistance, Excl Debt (Constant Prices, 2007 USD millions)" from the OECD Development Database on Aid from DAC Members (subset 2a). For GDP/capita we use the variable "Real GDP per capita (2005 constant price, Chain series)" from Penn World Table 6.3.

25. The method relies on three steps. First, we compute an average of the ordered violence variable for each country. Second, we define a new binary variable, as observations of the ordered variable above or below the country-specific averages computed in step one. Third, we estimate a conditional logit for the binary variable defined in step two. Building on Chamberlain (1980), Ferrer-i-Carbonell and Frijters (2004) show that this three-step procedure implements—in our context—an ordered logit with fixed country effects and country-specific thresholds.

TABLE I
BASIC RESULTS

Dependent Variable	(1) Ordered Variable	(2) Ordered Variable	(3) Ordered Variable	(4) Political Violence	(5) Political Violence	(6) Civil War	(7) Civil War	(8) Ordered Variable
Natural disaster	0.263**	0.317***	0.299***	0.278**	0.327***	0.370**	0.431***	0.263**
	(0.107)	(0.110)	(0.111)	(0.109)	(0.112)	(0.152)	(0.155)	(0.111)
Security Council member	−1.048***	−1.194***	−1.382***	−1.110***	−1.269***	−1.360**	−1.383**	−1.048**
	(0.399)	(0.417)	(0.456)	(0.412)	(0.430)	(0.545)	(0.547)	(0.413)
Security Council member in Cold War	1.275***	1.461***	1.657***	1.267***	1.465***	1.074*	1.105*	1.275**
	(0.439)	(0.458)	(0.495)	(0.453)	(0.472)	(0.633)	(0.635)	(0.504)
Natural disaster × cohesive institutions		−0.701*	−0.333		−0.618*		−1.233**	
		(0.374)	(0.318)		(0.376)		(0.595)	
Security Council member × cohesive institutions		1.975*	2.940***		2.186*			
		(1.173)	(1.123)		(1.178)			
Security Council member in Cold War × cohesive institutions		−2.577*	−3.379***		−2.746**			
		(1.375)	(1.247)		(1.381)			
Cohesive political institutions measure		High executive constraints 1950–2005	Parliamentary Democracy 1950–2005		High executive constraints 1950–2005		High executive constraints 1950–2005	
Estimation method	FE Ordered Logit	FE Ordered Logit	FE Ordered Logit	FE Logit	FE Logit	FE Logit	FE Logit	FE Ordered Logit
Significance of interactions (p-value)		0.61	0.49		0.66		0.17	
Observations	4251	4251	4251	4251	4251	2061	2061	4251
Number of countries	97	97	97	97	97	49	49	97

Notes. The time period covered is 1950 to 2006. For definitions of variables refer to the text. Standard errors are in parentheses: *significant at 10%; **significant at 5%; ***significant at 1%. Robust standard errors are in columns (1)–(7) with bootstrapped standard errors in column (8). The p-value refers to the significance of a test of the hypothesis that coeff.[natural disaster × cohesive institutions] = −coeff.[natural distaster] & coeff.[Security Council member × cohesive institutions] = −coeff.[Security Council member] & coeff.[Security Council member in the Cold War × cohesive institutions] = −coeff.[Security Council member in the Cold War], where coeff. is the estimated coefficient on the variable in question. The reduced sample size in columns (6) and (7) is due to all countries that never had a civil war during this period being dropped.

We are generally agnostic about the "right" sign for the Security Council membership variables. We expect this variable to perhaps reflect an accountability effect of temporarily being in the international spotlight. Our main interest is in the interaction with the Cold War period (in the third row). As stated, we hypothesize that the strategic geopolitical motives for giving aid (in the form of cash or military assistance) to Security Council members would have been much stronger in the Cold War period than after 1990. This is indeed what the results in column (1) of Table I suggest.

In columns (2)–(3), we show that these effects of natural disasters and Security Council membership are only found for countries with noncohesive political institutions. This claim is substantiated by interacting our three variables of interest with an indicator for cohesive political institutions, measured either by high incidence of strong executive constraints or parliamentary democracy (as detailed). If our exogenous variables have no effect under cohesive institutions, the coefficients for the interacted variables should be of the opposite sign and equal in absolute value to the coefficients on the noninteracted variables. Table I shows that the interaction coefficients do indeed have the opposite sign in every case. Moreover, for both our measures of cohesive institutions, we cannot reject the hypothesis of no correlation between the exogenous variables and political violence in countries with cohesive institutions: the p-values for these tests are reported at the bottom of the table. The results in these columns corroborate a key prediction of the theory.

It is reasonable to ask if these interaction effects really capture the effect of political institutions rather than just high income. To investigate that, we created a dummy variable that is equal to 1 if a country is in the top quarter (or top half) of the income per capita distribution in 1980. The correlations between this indicator of high income and the executive-constraints and parliamentary democracy measures of good institutions turn out not to be particularly high: 0.35 and 0.28, respectively (0.28 and 0.19, for the top half of income). When we add interactions of high income and shocks to the earlier specifications, all the results on the interactions with political institutions—both those above and those below—hold up qualitatively.[26]

26. The results are available from the authors on request.

In columns (4)–(7), we consider separately each of our predictable margins, namely, peace versus some violence (repression and civil war), and non-civil war (peace and repression) versus civil war. In each case, we estimate conditional logits that allow for country (and year) fixed effects. We report two specifications—one without and one with interaction terms for our executive-constraints measure of cohesive institutions. Columns (4)–(5) show that the earlier results are robust, with signs and magnitudes of the coefficients from the conditional logits being similar to those from the ordered logits. Again, we cannot reject the hypothesis that political violence in the cohesive institutions countries display no significant correlation with the exogenous variables. For the civil-war margin, only 49 countries have some time variation in the left-hand-side variable. We are unable to estimate an interaction effect with Security Council membership, since none of the cohesive institutions countries which have been on the Security Council have had a civil war during our time period. However, for the case of natural disasters, we cannot reject a zero effect for natural disasters on civil war in countries with cohesive political institutions.

These estimates square well with the predictions of our theory. The civil-war result is also consistent with the findings of Miguel, Satyanath, and Sergenti (2004) based on rainfall shocks rather than natural disasters, although here we have extended the sample from Africa to the world and widened the scope to include one-sided, in addition to two-sided, political violence. It is also consistent with the findings of Nel and Righarts (2008) who argue that natural disasters increase the risk of civil conflict, although our results are based exclusively on the within-country variation in the data (rather than the cross-sectional *cum* time-series variation).

Columns (1)–(7) all show nonadjusted standard errors. Since the estimation procedures are somewhat involved, the best alternative is probably to bootstrap (by country block) the standard errors. Whenever our bootstrapping procedure converges, it yields standard errors very similar to the nonadjusted standard errors.[27] Column (8) shows this by reporting bootstrapped standard errors for the same specification as in column (1). Reassuringly, the linear

27. The bootstrapping is nontrivial to perform due to the stepwise estimation (see the previous note) and the unbalanced panel, especially when the interaction effects in columns (2)–(3), (5), and (7) are included.

probability estimates in Table II rely entirely on standard errors that are robust to arbitrary forms of heterogeneity and serial correlation (Huber-White standard errors clustered at the country level).

Extended results. Table II looks at an alternative estimation method and also explores the mechanism at work in more detail.

The first four columns demonstrate that similar results are found when running the specifications in columns (4)–(7) of Table I with a conventional linear probability model with fixed effects. (Because we do not want to impose a strong cardinality assumption, we focus on the binary variables corresponding to the two margins investigated in Table I.) The standard errors in column (1), as in the whole of Table II, are robust to heteroskedasticity and clustered at the country level.

It is easy to give a direct quantitative interpretation of these estimates: having (at least) one natural disaster raises the probability of political violence by about 2.4 percentage points, and the probability of civil war by 2.9 percentage points. Security Council membership during the Cold War raises the probability of political violence by a whopping nine percentage points, compared to the post Cold War period. All of these effects appear quite large and consistent with the findings in Table I. The estimates of interaction effects with cohesive institutions, as measured by executive constraints, also display the same sign pattern as in Table I.

In columns (5)–(6), we investigate the potential mechanisms behind the reduced-form results that we have estimated so far. Specifically, we ask how our three exogenous variables affect two intermediate variables that the theory suggests might shape political violence—the logs of income per capita (for real wages) and aid disbursements.[28] In column (5) of Table II, we allow natural disasters and Security Council memberships to affect income per capita (allowing for income convergence by including the two-year lag of income per capita). The results show no significant correlation between these variables and income per capita. Although we cannot reject a negative effect of natural disasters on income per capita, it would be difficult to argue that the real wage is the main channel by which natural disasters affect the probability of conflict. In column (6), the dependent variable is instead (the log of)

28. Two recent studies of the relation between aid and civil conflict are de Ree and Nillesen (2009) and Nunn and Qian (2010).

TABLE II
EXTENDED RESULTS

Dependent Variable	(1) Political Violence	(2) Political Violence	(3) Civil War	(4) Civil War	(5) Log GDP per Capita	(6) Log Aid Disbursements
Natural disaster	0.024*	0.029*	0.029**	0.043***	−0.005	0.105**
	(0.013)	(0.017)	(0.013)	(0.016)	(0.003)	(0.043)
Security Council member	−0.066**	−0.092***	−0.051**	−0.053**	0.009	−0.269***
	(0.027)	(0.029)	(0.023)	(0.023)	(0.008)	(0.092)
Security Council member in Cold War	0.090**	0.129***	0.034	0.036	−0.004	0.434***
	(0.040)	(0.045)	(0.029)	(0.029)	(0.010)	(0.113)
Natural disaster × cohesive institutions		−0.024		−0.079***		
		(0.037)		(0.024)		
Security Council member × cohesive institutions		0.148***				
		(0.054)				
Security Council member in Cold War × cohesive institutions		−0.205***				
		(0.068)				
2-year lagged log GDP per capita					0.905***	
					(0.013)	
Observations	5880	5880	5880	5880	6300	5067
Number of countries	158	158	158	158	178	150
R-squared	0.030	0.031	0.056	0.059	0.914	0.136

Notes. The time period covered is 1950 to 2006. For definitions of variables refer to the text. Robust standard errors adjusted for clustering by country in parentheses (*significant at 10%; **significant at 5%; ***significant at 1%). The specification in all columns (1) through (6) is OLS.

aid disbursements. The estimates show that aid flows increase significantly with natural disasters, are higher during the Cold War when a country is on the Security Council, and are lower in the post-Cold War period. This sign pattern is identical to the effects of these variables on political violence.[29]

It is simple to compute the implied (semi)elasticity of political violence (p) with respect to aid, by observing that:

$$\frac{\partial p}{\partial \log(aid)} = \frac{\frac{\partial p}{\partial x}}{\frac{\partial \log(aid)}{\partial x}}.$$

Through this formula, the estimated coefficients in columns (1) and (6) give us three estimates of the elasticity of political violence to aid, which are remarkably similar—all in the range between 0.20 to 0.24. Quantitatively, a 10% increase in aid is therefore associated with an increase in the probability of violence by about two percentage points. These results are consistent with the recent results on aid and civil conflict presented by Nunn and Qian (2010).[30]

Taken together, we believe that the empirical estimates presented in Tables I–II are consistent with the theoretical predictions derived in Section II and operationalized in Section III.[31]

29. We have also interacted these shocks with our institutional measure (available from the authors on request). For natural resource shocks, we find that they (significantly) increase aid in countries with weak institutions, but (significantly) reduce aid in countries with strong institutions. However, Security Council membership has a (significantly) much stronger effect on aid, both during and after the Cold War, in countries with strong institutions than in countries with weak institutions. The latter result suggests that it is really the difference in institutions that drive the results in Table I, rather than a different response to shocks.

30. Nunn and Qian (2010) use weather shocks in the U.S. wheat belt to instrument for U.S. food aid across the world. Their results and those in this article are at odds with de Ree and Nillesen (2009) who study civil conflicts in Sub-Saharan Africa. They use shocks to GDP per capita in the United States and a few other donor countries to instrument for Overseas Development Assistance in Sub-Saharan Africa. Their dependent variable is also different. they find significant effects only when they study the persistence and onset of civil war.

31. If we estimate an instrumental variables specification where income per capita and aid are instrumented with our exogenous variables, then we find a positive and significant effects of aid dispersement on political violence as well as civil war. Moreover, the estimates are close in magnitude to the aid-to-violence elasticities computed from our earlier estimates. However, the assumption that the exogenous variables enter only via measured aid and GDP per capita is too doubtful to push these results.

V. Final Remarks

This article takes some steps toward integrating two different strands of research on political violence, developing a theoretical model to analyze the common roots of repression and civil war. Under specific assumptions about the conflict technology, we show that peace, repression (one-sided violence), and civil war (two-sided violence) become ordered states depending on a common underlying latent variable which is shifted by shocks to the value of public goods, wages, aid, and resource rents. But these effects only emerge when political institutions provide insufficient checks and balances or enough protection for those excluded from power.

The article also bridges the gap between theoretical modeling and econometric testing. Under specific assumptions on what can be observed, our model's predictions can be taken to the data by estimating either a fixed-effects ordered logit, or the conditional probability of transition from peace to violence or from non-civil war to civil war.

Our empirical strategy makes use of two sources of, arguably, exogenous variation affecting violence, which make sense in terms of the underlying theory: natural disasters (affecting real wages and aid flows) and membership of the Security Council (affecting aid flows). The empirical results are consistent with the theoretical predictions in that these variables indeed alter the likelihood of government repression, as well as civil war, in line with our theoretical priors. However, this is the case only if checks and balances are weak and/or there is weak minority representation. Inspecting the mechanism, we find that variations in foreign aid seem to be consistent with the within-country variations in political violence that we explain.

These findings resonate with previous work that emphasizes the role of institutions, economic development and natural resources in shaping civil conflict, or political violence more generally. However, much work remains to complete the agenda of interpreting empirical results on violence through the lens of well-specified theoretical models. One helpful, but limiting, feature of the current model is the symmetry between incumbent and opposition groups. The model can be extended to incorporate income inequality via heterogeneity in wage rates. Groups might also differ in their weighting of national interests (national public goods) against group-specific interests (transfers), which could offer a way to model ethnic, cultural, or religious tensions. The way that heterogeneity impacts on political violence is more subtle than is often claimed based on intuitive reasoning.

Our empirical analysis of the *incidence* of violence has not really engaged with the distinction between *onset* and *duration* of violence, which plays an important role in the empirical civil war literature. To make further theoretical progress on this issue would require specifying an underlying source of state dependence. We could get a genuinely dynamic model by also introducing asymmetries between the groups. The state variable would then be the group in power, making the equilibrium in any given period state-dependent. This would naturally lead to an empirical model where political violence and political turnover are jointly determined. Another possibility would be to introduce an economic state variable such as land or capital, with conflict in one period cutting this state variable in the next. The implied dynamics of the real wage would naturally imply some duration dependence in conflict.

More generally, it would be interesting to study—theoretically and empirically—the two-way links between political violence and economic development. This is a difficult issue, but a start is made in Besley and Persson (2010, 2011), who use the framework in Besley and Persson (2009b) to study interactions between political conflict and the building of state capacity where state development goes hand in hand with economic development.

Appendix

Proof of Proposition 2

Proof. To simplify the notation, the proof leaves out the dependence of γ on parameter vector ξ. The first-order conditions for the problems faced by L^I and L^O are:

$$-\gamma_I\left(L^O, L^I\right)\left[Z - L^I\right](1-2\theta) - \left[1-\theta-\gamma\left(L^O, L^I\right)(1-2\theta)\right] = 0$$

and

$$\left[2\gamma_O\left(L^O, L^I\right)(1-2\theta)\left[Z-L^I\right]-1\right]L^O = 0 \text{ for } L^O < X$$

$$\left[2\gamma_O\left(V, L^I\right)(1-2\theta)\left[Z-L^I\right]-1\right] \geq 0 \text{ otherwise.}$$

Observe that with $\gamma \in (0,1)$ we can ignore the upper bound $L^I = Z$.

First, we show that at any interior solution, resources devoted to fighting by both groups is increasing in Z. To see this, note that differentiating and using the first-order conditions yields:

(13)

$$\begin{bmatrix} -\frac{\gamma_{II}}{2\gamma_O} + 2\gamma_I\left(1-2\theta\right) & \gamma_O\left(1-2\theta\right) - \frac{\gamma_{IO}}{2\gamma_O} \\ \frac{\gamma_{IO}}{\gamma_O} - 2\gamma_O(1-2\theta) & \frac{\gamma_{OO}}{\gamma_O} \end{bmatrix} \begin{bmatrix} dL^I \\ dL^O \end{bmatrix} = \begin{bmatrix} \gamma_I\left(1-2\theta\right) \\ -2\gamma_O\left(1-2\theta\right) \end{bmatrix} dZ.$$

Define $\Omega = \frac{\gamma_{OO}}{\gamma_O}\left[-\frac{\gamma_{II}}{2\gamma_O} + 2\gamma_I\left(1-2\theta\right)\right] + 2\left[\frac{\gamma_{IO}}{2\gamma_O} - \gamma_O(1-2\theta)\right]^2 > 0.$

Solving Equation 13 using Cramer's rule yields:

$$\frac{dL^I}{dZ} = \frac{\left(1-2\theta\right)\left[\left(\frac{\gamma_i\gamma_{OO}}{\gamma_O} - \gamma_{IO}\right) + 2\left(\gamma_O\right)^2\left(1-2\theta\right)\right]}{\Omega} > 0$$

and

$$\frac{dL^O}{dZ} = \frac{\left(1-2\theta\right)\left[\left(\gamma_{II} - \frac{\gamma_I\gamma_{IO}}{\gamma_O}\right) - 2\gamma_I\gamma_O\left(1-2\theta\right)\right]}{\Omega} > 0\,,$$

where we have used both parts of Assumption 1c.

We now derive two trigger points for violence. Define $\hat{L}\left(Z\right)$ from

$$\begin{aligned} -\gamma_I\left(0,\hat{L}\left(Z\right)\right)\left(1-2\theta\right)\left(Z-\hat{L}\left(Z\right)\right) \\ -1+\theta+\gamma(0,\hat{L}\left(Z\right))(1-2\theta) \leq 0 \\ \text{c.s. } \hat{L}\left(Z\right) \geq 0. \end{aligned}$$

It is simple to check that this is an increasing function of Z under Assumption 1a. Clearly with $L^O = 0, L^I = \hat{L}\left(Z\right)$. We can define $Z^I(\theta)$ from $\hat{L}\left(Z\right) = 0$, that is,

$$Z^I(\theta) = \frac{\frac{1-\theta}{1-2\theta} - \gamma\left(0,0\right)}{-\gamma_I\left(0,0\right)}.$$

Next, define $Z^O(\theta)$ implicitly from

$$2\gamma_O\left(0,\hat{L}(Z^O(\theta)\right)\left(1-2\theta\right)\left(Z^O(\theta) - \hat{L}\left(Z^O(\theta)\right)\right) = 1.$$

The expression for $\frac{dL^O}{dZ}$ implies that for $Z \geq Z^O$, we must have $L^O > 0$.

As the next step, we prove that $Z^O(\theta) > Z^I(\theta)$. Suppose not, then

$$\gamma_O(0,0)(1-2\theta)Z^O(\theta) = \frac{1}{2}.$$

If so,

$$Z^O(\theta) = \frac{1}{\gamma_O(0,0)(1-2\theta)} \leq Z^I = \frac{\frac{1-\theta}{1-2\theta} - \gamma(0,0)}{-\gamma_I(0,0)},$$

or

$$\frac{-\gamma_I(0,0)}{\gamma_O(0,0)} < 2(1-\theta)\left(\frac{1-\theta}{1-2\theta} - \gamma(0,0)\right) < 2\left[1-\gamma(0,0)\right],$$

which contradicts Assumption 1b for all values of θ.

Finally, it is easy to see from the explicit definition that $Z^I(\theta)$ is an increasing function. Using the implicit definition of $Z^O(\theta)$, and the fact that $\hat{L}\left(Z^O(\theta)\right)$ is increasing, it follows that this function is increasing as well. This concludes the proof of the proposition. ■

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