The Logic of Political Violence*

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Abstract

Political violence is an everyday occurence in many weakly institutionalized polities. This paper offers a unified approach for studying political violence whether it emerges as repression or civil conflict. We formulate a model where an incumbent or opposition can choose to use violence as a means of acquiring or maintaining power. We study the institutional and economic factors that determine the use of one-sided or two-sided violence (repression or civil war). The model gives way naturally to a hierarchy of violence from repression to civil war, which forms the basis for our empirical approach. Accordingly, we construct an ordered variable to explore the empirical determinants of violence. Two robust factors emerge from the data. First, economic shocks are robustly correlated with the use of violence. Second, the relationships are heterogeneous depending on political institutions. All our results are based on exploiting only the within-country variation in violence.

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1 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 Uppsala/PRIO data set, is over 10%, with a peak of more than 15% in the early 1990s. The upper left part of Figure 1 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 the year 1980. Clearly, civil conflicts are disproportionately concentrated to the poor countries of the world. The cumulated death toll of these conflicts is now approaching 20 million people.[1]

A key feature of civil war is two-sided violence, typically between an insurgent and the government. However, many citizens suffer consequences of one-sided political violence due to government repression manifested in a variety of infringements of human rights. The Banks (2005) data set reports a stark form of repression viz. purges – i.e., the removal, by jailing or assassination, of opponents considered undesirable by the incumbent government. Since 1950, more than 7% of all country-years are associated with purges, in the absence of outright civil war. The lower left graph in Figure 1 shows the trendwise worldwide development of repression, as measured by the purges. Interestingly, up to the early 1990s, this series is almost a mirror image of the prevalence of civil war up in the graph above. If we plot the prevalence of repression by country against the level of a country's GDP in 1980, a striking observation is that repression is most common in richer countries compared to those where civil war is prevalent.

Our paper argues that an important common element in political violence in both of its forms is weak political institutions. We provide a framework that allows us to study one-sided and two-sided violence together, which leads to an empirical strategy for testing its determinants. The research we present builds on a large existing literature that has progressed 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 has progressed from mainly cross-sectional inference using country level data to panel data studies, which exploit within-country variation – see the survey by

[1] See Lacina and Gledtsch (2005).

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 of both of these strands of work has been on exploring empirical regularities, searching in many cases for credibly exogenous sources of variation. Links between theoretical models of conflict and violence are limited, such that both Blattman and Miguel (2009) and Davenport (2007) lament the fact that few of the empirical findings forge links between the theory and data.[2] In this paper, we argue that a theoretical framework provides a natural join between forms of political violence, seeing civil war and repression as manifestations of similar pathologies.[3]

In Section 2, we put forward a simple "canonical" model, where an incumbent government and an opposition group each can make investments in political violence. The resulting conflict game is embedded in a policy setting, where the ruling group in each period obtains control over 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. Moreover, we identify specific conditions on the conflict technology, under which these three conflict states are *ordered* in a latent variable, which summarizes some 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. We derive additional comparative statics with regard to minority protection (representation) in political institutions, and parameters of the conflict technology. Our theoretical results are summarized in two propositions and a number of corollaries.

In Section 3, we discuss how the predictions of the model can be used to 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 from 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 the within-country variation in the drivers of conflict – in our case,

[2] There are certainly exceptions, however such as Dube and Vargas (2008), who build explicitly on the theoretical framwork developed by Dal Bo and Dal Bo (2006). See also Fearon (2008)

[3] In a short previous paper, Besley and Persson (2009a), we brought out some of these ides in a simple linear example.

as suggested by the theory. Moreover, the theoretical framework gives way naturally to an ordered logit specification, with fixed country and year effects, for the three states: peace, repression and civil war.

Section 4 describes the data we use to measure political violence in the forms of civil war and government repression. We exploit three different sources of, arguably, exogenous variation: natural disasters (for shocks to wages), membership in the UN Security Council during the Cold War period (for shocks to aid flows), and fluctuations in world commodity prices (for shocks to resource rents). The estimation results provide quite strong empirical support for the specific model predictions. Thus, lower wages and higher aid flows or resource rents both raise the probability of political violence – i.e., civil war or repression – but this only happens under non-inclusive political institutions. In all cases, we look for heterogeneity according to weak and strong political institutions identified from the data as detailed below. In line with the priors from the theoretical model, it is the combination of shocks coupled with weak institutions that drive the empirical findings.

Overall, the paper provides a step towards integrating two separate literatures both from a theoretical and empirical point of view. While we do not provide any general literature review, we will relate our approach more precisely to the existing literature, as we go along in the sections to follow. An Appendix collects the proofs of some theoretical results. Section 5 concludes the paper.

2 Theoretical Framework

Our aim is to build a simple and tractable model that serves as a useful guide to how observable economic and political factors might 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 to (post-conflict) strategies. The key friction in our model is of the second type: an inability of any prospective government to credibly offer post-conflict transfers, and the inability of potential insurgents to commit not to use their capacity to engage in conflict.

There are two groups: A and B. Each group makes up one half of the population. Time is infinite and denoted by $t = 1, ...$, although we will drop the time index in this theoretical section. One generation is alive at each date and is labelled according to the date at which it lives. There are no state variables in the model. 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 $O \in \{A, B\}$. The incumbent group can make an investment in violence – think about mounting an army – denoted by L^I, which is financed out of the public purse. Power can be transferred by peaceful means. But the opposition can also invest in violence – think about mounting an insurgency – with armed forces L^O and try to take over the government. The conflict technology is discussed below. The winner, whether power is transferred peacefully or through armed conflict, 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 e.g. 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 $T^{I'}$ and the opposition $T^{O'}$. 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 below.

Individual incomes and utility Individuals supply labor in a common labor market to earn an exogenous wage w. Like revenues, wages are stochastic and (iid) distributed on finite support: $w \in [w_L, w_H]$. We assume that individuals have 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

$$R - \sum_{J \in \{I', O'\}} \frac{T^J}{2} - G - wL^I \geq 0 \ , \tag{2}$$

where L^I denotes the size of the army chosen by the incumbent.

Institutions As mentioned above, power can be transferred between groups according to democratic principles, or by a violent conflict in which groups may raise armed forces L^J to fight. The probability that group O wins power and becomes the new incumbent I' is

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

which depends on the resources devoted to fighting and some parameter vector, $\boldsymbol{\xi}$. The relative probabilities of winning are thus determined by a contest function, as in much of the existing theoretical literature on domestic conflict (Dixit, 1987 and Skaperdas, 1996 survey the use of contest functions). We assume this function $\gamma(\cdot)$ is increasing in its first argument and decreasing in the second. In this notation, $\gamma\left(0, 0; \boldsymbol{\xi}\right)$ is the probability of a peaceful transition of power between the groups. Below, we make specific assumptions on the functional form of (3).

Each group (when in opposition) has the power to tax/conscript its own citizens to finance a private militia in order to mount an insurgency. We denote this capacity by X so $L^{O_{s-1}} \leq X$ 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 a politician must give $\sigma \in [0,1]$ to the the opposition group, when it makes a transfer of 1 to its own group implying that $T^{O'} = \sigma T^{I'}$. Given this assumption, we use the government budget constraint (assuming that it holds with equality) to obtain:

$$T^{I'} = 2\left(1-\theta\right)\left[R - G - wL^I\right] \ , \tag{4}$$

where $\theta = \frac{\sigma}{1+\sigma} \in [0, 1/2]$. Throughout, we interpret a higher value of the opposition's share of transfers, θ, as reflecting more representative, or consensual, 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 $\theta = 1/2$, then transfers are shared equally across the two groups. Thus, we can think of θ as an institutionalized ability of making commitments not to expropriate the opposition; thus, θ closer to (further from) one half is 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 $1-\gamma\left(L^{O},L^{I}\right)$.
4. The winning group becomes the new incumbent I' and determines policies, i.e., spending on transfers $\left\{T^{J}\right\}_{J\in\{I',O'\}}$ and public goods G.
5. Payoffs are realized, consumption takes place, and the currently alive generation dies.

We next solve the model by working backwards to derive a sub-game perfect equilibrium.

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

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

Define $\widehat{G}\left(z\right)$ by

$$H_{G}\left(\widehat{G}\left(z\right)\right)=\frac{1}{z}\ .$$

We record the policy solution as:

Lemma 1 *For given (α,w,R), public goods are provided as*:

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

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. In the special case when $\theta=1/2$, we get the same amount of spending on public goods as the amount that would be chosen by a Utilitarian planner, namely $\widehat{G}(\alpha)$. 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 process of conflict, looking for an equilibrium in which the opposition decides whether to mount an insurgency and the incumbent government chooses how to respond. As we show below, the equilibrium has three possible regimes. In the first, no resources are invested in violence by either side, i.e. 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:

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

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

For the opposition group, we have

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

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

These payoff functions expose a key asymmetry in the model 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. For the opposition, any insurgency must be financed 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 the 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 and take advantage of the monopoly on allocating government revenue.

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

We characterize a Nash equilibrium in pure strategies where:

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

for the incumbent and

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

for the opposition.

We first state a simple result:

Proposition 1 *Let $\left(\hat{L}^I, \hat{L}^O\right)$ be a Nash equilibrium of the conflict game. Then, as $\theta \to 1/2$, there is always peace.*

Proof. When $\theta \to 1/2$, 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 hold on to power or depose the existing government. This gives a simple account for why we predominantly observe political violence in countries with weak political institutions.

To study the Nash equilibrium when institutions do not make a country conflict proof, we make the following assumption on the properties of the conflict technology (let γ_O and γ_I denote derivatives with regard to the first and second argument, respectively)

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

a. $\gamma \in (0,1)$,

b. $\gamma_O > 0, \gamma_I < 0$, $\gamma_{OO} < 0$, $\gamma_{II} > 0$,

c. $\frac{-\gamma_I(0,0;\boldsymbol{\xi})}{\gamma_O(0,0;\boldsymbol{\xi})} \geq 2\left[1-\gamma\left(0,0;\boldsymbol{\xi}\right)\right]$, *and*

d. $\frac{\gamma_I\gamma_{OO}}{\gamma_O} \geq \gamma_{IO} \geq \frac{\gamma_O\gamma_{II}}{\gamma_I}$.

Condition **a** just says that neither group can ever be fully certain of holding power, whereas **b** says that fighting always has positive returns for both groups, albeit at a decreasing rate. The property in **c** 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, **d** restricts the extent of any strategic complementarities or substitutabilities in the conflict technology.

Assumption 1 is satisfied by a number of reasonable and widely used contest functions. For example, it holds in 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 $\xi \geq 1$. It also holds in the semi-linear formulation:

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

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

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

Proposition 2 *Let* $\left(\hat{L}^I, \hat{L}^O\right)$ *be a Nash equilibrium of the conflict game. Then, if Assumption 1 holds and* $\theta < 1/2$, *there exist two thresholds* $Z^I(\theta;\xi)$ *and* $Z^O(\theta;\xi)$ *with* $Z^I_\theta, Z^O_\theta < 0$ *and*

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

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

Assumption 1d essentially guarantees that both incumbent and opposition have propensities to fight that are increasing in the size of the prize, as 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 changes in the marginal return to fighting are not strongly affected by the fighting decisions of the other group, placing bounds on γ_{IO}, not allowing a too strong positive or negative cross-partial. [4]

[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 1d. This amounts to saying that the conflict

Implications Our results have some striking empirical implications. They give the logic of political violence expressed as a function of latent variable Z. More precisely, our theory predicts an *ordering* of the three states peace, repression, and civil war, in Z. This ordering is particularly interesting against the backdrop of Figure 1 above, which indeed 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 world.

The Z variable summarizes several important determinants of violence, which we now bring out in a set of corollaries. We state these corollaries in terms of likelihoods, implicitly assuming that some factors are not only uncertain but also not observed by an outside analyst. A more precise formulation of the empirical predictions – along precisely these lines – will be made in the next section.

Corollary 1 *Higher wages, w, reduce the likelihood that an economy will experience political violence, i.e., be in repression or civil war.*

The result follows by observing that w is the denominator of the expression for Z. In other words, given the distributions of α and R, when w is higher the whole distribution of Z shifts to the left. Thus 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 the state of repression becomes more or less likely depends on relative densities (in the p.d.f. of Z).

Of course, this result reflects that higher wages raise the opportunity cost of fighting and hence reduce the relative 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). Here, we see that the result extends to political violence more generally.

Corollary 2 *Higher natural resource rents. or any other exogenous forms of income such as aid, a higher R, increase the likelihood that an economy will be in repression or civil war.*

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technology is quasi-concave, i.e. has level sets that are convex in $\left(L^O, L^I\right)$ space. This makes total spending on conflict by both parties monotonic in Z, but not necessarily the spending by each group is monotonic. 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.

The corollary follows directly from noting that Z depends directly on the level of natural resource rents or exogenous income from any other sources that accrue to government. 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.*

This follows from the fact that an increase in α increases G and hence reduces Z. To the best of our knowledge, this specific prediction of our model is new, even to the formal modeling of civil war, since the modeling of conflict is typically not embedded in an explicit public finance context.

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 changes in the parameters of the conflict technology $\boldsymbol{\xi}$, 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 (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 both $Z^O(\theta; \boldsymbol{\xi})$ and $Z^I(\theta; \boldsymbol{\xi})$ are decreasing functions of θ. Intuitively, more inclusive institutions make control of the state less valuable, and thus shift up the points 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 and 2 also have the joint implication that Corollaries 1-3 should only be expected to 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.

3 From Theory to Econometric Testing

In this section, we discuss how our proposed theory can inform an empirical study of the determinants of political violence. Although our model is extremely simple, it does give a transparent set of predictions on how parameters of the economy and the polity affect the incidence of conflict. A clear advantage of beginning from the theory is that we can explicitly clarify and evaluate the assumptions made on the way to empirical testing of the model predictions. Specifically, we must take a stance on which variables and parameters are measurable in the data – i.e., which are observable and which are unobservable – as well as which variables and parameters to treat as fixed (at the country level) and which to treat as time varying.

Measurement, observability and likelihoods Our data are in panel form for countries and years from 1950 onwards. Hence, consider country c at date t. Below, 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 below.

However, we will not be able to measure variations in public goods, as induced by time-varying parameter $\alpha_{c,t}$. 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 $[\widehat{G}(\frac{\alpha_L}{2(1-\theta_c)}) - G_c, \widehat{G}(\frac{\alpha_H}{2(1-\theta_c)}) - G_c]$. As for the other parameters of the model, we will treat them as constant over time. Finally, while we will be able to observe proxies for the inclusiveness of political institutions, θ_c, we will not readily observe the parameters of the conflict technology, $\boldsymbol{\xi}_c$.

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

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

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

$$F^c(R_{c,t} - Z^O(\theta_c; \boldsymbol{\xi}_c)\, w_{c,t} - G_c) \ . \tag{8}$$

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 $\frac{1}{2}$.[5]

By similar reasoning, the likelihood of observing peace is

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

while the likelihood of observing repression is

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

As explained in Section 2, the theory gives us distinct predictions regarding how changes in $R_{c,t}$ and $w_{c,t}$ shift the distribution of index variable $Z_{c,t}$ and thereby change 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 of how such changes affect 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 (8)-(10) is that these expressions define the relative probabilities of the three ordered states of violence. This strongly suggest that the most direct and straightforward way of confronting the theory with data would be to estimate a fixed-effect ordered logit driven by variables that shift the country-specific distribution of $Z_{c,t}$ given the country-specific thresholds $Z^I\left(\theta_c; \boldsymbol{\xi}_c\right)$ and $Z^O\left(\theta_c; \boldsymbol{\xi}_c\right)$.

Cross country vs. 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, estimate the probability of observing civil war or repression, respectively, using cross-sectional data sets. Expressions (8)-(10) illustrate clearly why this may not be such a good idea. Cross-sectional data will 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, since we run a large risk of confounding the cross-country variation in these variables with cross-country variation in the unobserved parameters G_c and $\boldsymbol{\xi}_c$, something which could seriously bias and invalidate the estimates.

[5] Formally, as θ approaches 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.

It will be more rewarding to exploit within-country variation in panel data, as in the cross-country panel studies of civil war in Africa by Miguel, Satayanath 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 Columbia. For instance, estimating a specification for the likelihood of observing civil war, with fixed country effects, is equivalent to evaluating

$$F^c(R_{c,t} - Z^O\left(\theta_c; \boldsymbol{\xi}_c\right) w_{c,t} - G_c) \; - E\{F^c(R_{c,t} - Z^O\left(\theta_c; \boldsymbol{\xi}_c\right) w_{c,t} - G_c) \; \} \; , \tag{11}$$

i.e., 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. The impact of their average values as well as the time-invariant parameters in each country will be absorbed by the country fixed effect.

Given the important and irregular time trends in the prevalence of civil war and repression in Figure 1, 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 will then be picked up in a flexible (non-parametric) fashion, and we will only be using the country-specific yearly variation relative to world year averages for identification.

To be more precise about the specification, and take into account that the predictions regarding the effect of 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 1/2) in country c in the period of our data, and equal to zero otherwise. We then model the index function in (11) as:

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

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 which reflect changes in $R_{c,t}$ and $w_{c,t}$. Since the theory predicts that the parameter of interest, $b\left(\Theta_c\right)$, is heterogeneous with respect to Θ_c, we will estimate the model separately for strongly and weakly institutionalized countries.

4 Data and Results

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

Data: political violence and political institutions A large body of literature looks at the determinants of civil war.[6] In this paper, we will mainly use the UCDP/PRIO civil-war incidence measure, starting in 1950. This measure takes a value of 1 if a given country in a given year is involved in a violent conflict, which claims a (cumulated) death toll of more than 1000 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.[7]

To measure repression, we use a measure from Banks (2005), which counts up purges: systematic murders and eliminations of political opponents by incumbent regimes. We create an indicator which is equal to one in any year when purges exceed zero. Here, we use the data from 1950 onwards in our ordered logits. 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.[8]

Based on these two measures, we construct our ordered variable of political violence. Specifically, we assign a value of 0 to the state of peace, a value of 1 to the state of repression in the absence of civil war, and a value of 2 to the state of civil war.[9]

Are the three states peace, repression and civil war naturally ordered in the data, as in the theory? For income per capita, the answer is a clear-cut yes. Peaceful countries have an average GDP per capita of $4,365, repressing

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

[7] An alternative measure is available in the Correlates of War (COW) data base, which only runs up to 1997. However, given that one of our independent variables relies on the experience before and after the cold war, the COW variable would be quite limiting, giving us only eight, as opposed to sixteen, annual observations in the post-cold war era.

[8] An alternative would be to use the commonly used Political Terror Scale based on the reports on human-rights violations by the US State Department and Amnestsy 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 Yanagisawa (2009), security council membership during the cold war period may have affected the way the US State department reported on human rights in US allied and non-allied countries.

[9] To be precise, we begin from two underlying variables: civil wars as coded in the UCDP/PRIO data set and the purges variable in the Banks (2005) data set. We construct a binary variable based on the latter denoting whether there are some purges in a country at a given date. Since 1950, there are 4841 observations where there is neither civil war or government purges. There are 90 observations where there is both a civil war and the government is using purges. There are 714 observations where there are civil wars but no purges and 425 observations where there are purges but no civil war. This yields 1229 observations where there is some violence and 804 where there is a civil war.

countries are poorer with \$2,503 per capita, while the countries in civil war are the poorest with average incomes of \$1,789.

We construct two indicator variables to capture strong political institutions, corresponding to Θ_c in Section 3. Our core measure is based on the assessment of executive constraints in the Polity IV data. We believe this variable best captures the thrust of θ in our theory. When coding it, we adopt quite a conservative approach. We first evaluate the pre-sample evidence, measuring the fraction of years for which a country had the highest score (of 7) 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 good political institutions, $\Theta_c = 1$, only if the fraction in the pre-sample period is above zero, *and* the fraction the sample period is greater than 60%. This is a very conservative definition and classifies less than 18% of countries as having good institutions.

Using this variable, we uncover a striking regularity across political regimes: 19.6% of the country-years with peace occur in countries with good institutions, while the number falls to 7.3% for repression, and 5.8% for civil wars.

As a robustness check, we use an alternative measure of political institutions based on the prevalence of parliamentary democracy. While high executive constraints are associated with stiffer checks and balances on the government, this measure is intended to capture larger representativeness.[10] We define it analogously to the core measure, namely as the result of having had a positive prevalence of parliamentary democracy before 1950, and a minimum prevalence of 60% in between 1950 and 2005.

Data: Z-shocks In order to test the specific model predictions with the specification in (12), we still need credibly exogenous variation in the time-varying regressors $\widetilde{Z}_{c,t}$. We use three variables for this purpose. The first, which we expect primarily to affect the wage $w_{c,t}$, is a measure of natural disasters, is constructed from the EM-DAT data set. 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. We then create a binary indicator variable, set equal to one if a country experiences any such event. Consistently with our hypothesis that this variable reflects negative shocks to the wage, we find that a country-year with at least one natural

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

disaster is associated with a 2.5% reduction in income per capita. Of course, a natural disaster may also trigger international aid flows. In terms of our theory, this would correspond to a positive shock to $R_{c,t}$, which would affect 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 U.N. security council (for the non-permanent members). We expect membership to raise a country's geopolitical importance and therefore its susceptibility to receive international aid, corresponding to positive shocks to $R_{c,t}$. Indeed, Kuziemko and Werker (2006) find that US 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 the regime 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 stronger in the period before 1990, because of the stronger geopolitical tensions during the cold war.[11]

Finally, we exploit shocks to commodity prices in world markets to generate exogenous time variation in resource rents, corresponding to changes in $R_{c,t}$.[12] Using trade volume data from the NBER-UN Trade data set, and international price data for about 45 commodities (minerals as well as agricultural goods) from UNCTAD, we construct country-specific export price indexes.[13] The price index for each country has fixed weights, computed as the share of exports of each commodity in the country's GDP in a given base year (1980). We do not include oil in the export price index. Since oil price fluctuations are particularly large, and oil figures prominently in oil produc-

[11] Possibly, cold-war security-council membership may affect conflict through a different channel, namely the provision of military aid raising the government capability to fight. In the semi-linear conflict model mentioned in Section 2, 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 points) 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.

[12] A positive price shock may also raise the real wage $w_{c,t}$. thus driving the likelihood of political violence in the opposite direction. Such an indirect "Dutch disease" effect is unlikely to dominate the direct effect on the likelihood of conflict of higher reource rents $R_{c,t}$.

[13] The method that we follow is similar to Deaton and Miller (1996).

ing countries export basket, these fluctuations are hard to distinguish from a non–parametric trend and hence from the time dummies in our favored specification. Other commodity prices induce a much more heterogenous pattern of shocks across countries.

Basic results We present the empirical results in three tables. We start in Table 1 with a very conservative specification. This only includes the natural disaster indicator as a measure of wage shocks plus country and year fixed effects. The spartan specification allows us to contrast some different estimation methods. As we have already discussed, the natural treatment of our ordered variable given the theory, is to estimate an ordered logit. But before we do this, we follow the other approach discussed in Section 3, namely to consider each of our predictable margins separately, 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 fixed effects. These specifications appear in columns (1) through (6) of Table 1, where we also split the sample according to the strength of political institutions, as our theory tells us to do.

Column (1) shows that the negative wage shock associated with (at least) one natural disaster in a particular year significantly raises the probability of some political violence (repression or civil war). The effect is precisely estimated and its magnitude is non-trivial: the point estimate corresponds to just over 4 percentage points higher probability of observing violence, given a sample average of about 17%. As columns (2) and (3) show, this effect is only present in countries with low executive constraints, whereas we do not find a significant impact of wage shocks in countries with good political institutions. Columns (4) to (6) show that negative wage shocks also operate on the other predicted margin. Specifically, lower wages raise the likelihood of observing civil war, but only in countries with weak constraints on the executive. Again, the estimated effect is quantitatively significant, as a natural disaster raises the probability of civil war by just below 4 percentage points.

These estimates square well with the predictions of our theory. The civil war result is consistent with the findings of Miguel, Satyanath and Sergenti (2004) based on rainfall shocks rather than natural disasters, although here we extend the sample from Africa to the world and widen the scope to include one-sided political violence.

In columns (7) through (9), we present results for a fixed-effects ordered logit. This is not a standard estimation method, but we implement it in a way suggested by Ferrer-i-Carbonell and Frijters (2004).[14] As the results show, a negative wage shock raises the probability of repression and civil war. Again, the estimated effect comes only from the countries with weak political institutions.

Note that the fixed effects in the conditional and ordered logits throw out all countries that do not have any variation in violence, as the country dummies perfectly predict a constant violence state. For example, we have a total of 158 countries with violence and natural disasters data, while column (1) only uses 98 of them. The consequence of this shrinkage is particularly stark for the countries with good political institutions: because many of them are peaceful throughout, we are left with only 8 countries with some violence of any type, and only 2 with some civil war. These low numbers are entirely consistent with the theory: any positive shock to latent variable $Z_{c,t}$ has to be very large to push the country across the violence thresholds $Z^I\left(\theta_c;\boldsymbol{\xi}_c\right)$ and $Z^O\left(\theta_c;\boldsymbol{\xi}_c\right)$ in the case when θ_c is close to $\frac{1}{2}$.

Finally, all the specifications in Table 1 are based on non-adjusted standard errors. We have also run the same specifications with bootstrapped standard errors, and the results are virtually identical.

Extended results Table 2 broadens the scope by incorporating our additional variables for Z-shocks. Columns (1) to (3) present the results from fixed-effects ordered logits, on the whole sample and the two subsamples, with natural disasters plus our two security-council membership variables which are basically available for all countries in our panel. For the natural disasters variable the results are very similar to those in Table 1. Security-council membership only affects violence in the countries with bad political institutions. The general effect (in the second row) of membership seems to be negative. We are agnostic about this sign, which may 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).

[14] 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.

As stated above, our hypothesis is 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. In that interpretation, higher foreign aid associated with membership during the cold war significantly raises the probability of repression or civil war, but only in countries with weak political institutions.

The other half of Table 2 instead combines the security-council membership variables with our country-specific index for the prices of exported commodities. (Because the price index and natural disasters data are available for partly non-overlapping samples, we do not add together all the three variables in the same specification.) The estimates for security-council membership are very similar to those in the first half of Table 2. In the full sample (column (4)), higher world prices of exported commodities, corresponding to higher resource rents in our model, do indeed significantly increase the probability of repression or civil war. But this effect is due to the variation in countries with low executive constraints; in fact, the coefficient with high executive constraint is significant with the opposite sign.

When we rerun the specifications in Table 2 with bootstrapped standard errors, the estimates in the three first columns come out the same, whereas the estimates of the export price effects in the last three columns are considerably less precise. Presumably, this reflects the fact that our export price indexes display a high degree of serial correlation, whereas this is not true for our other two independent variables.[15]

The first four columns in Table 3 check that we get similar results when the two specifications in Table 2 are run with a conventional fixed-effect estimator, corresponding to a linear probability model. Since we do not want to impose a strong cardinality assumption, we focus on the binary variables corresponding to the two margins investigated in Table 1 and do not show any results for the ordered violence variable. On the whole the estimates are consistent with the earlier ones. Interestingly, whereas natural disasters and export price booms seem to operate on both margins, security-council membership during the cold war appears to operate mostly on the political-violence margin.

These estimates are easier to give a direct quantitative interpretation than the non-linear estimates. Thus, having (at least) one natural disaster

[15] In fact, we cannot reject that the export price indexes (on average) in the panel follow a random walk.

increases the probability of political violence by about 2.5 percentage points, and the probability of civil war by 3 percentage points. Having a hike of export prices by one standard deviation raises the probability of political violence by almost 6 percentage points (0.034·1.8), and the probability of civil war by almost 5 percentage points. Security-council membership during the cold war raises the probability of political violence by a whopping 8 percentage points, when compared to the post cold-war period, and 3 percentage points in absolute terms. All these effects appear quite large and consistent with the findings in Table 1.

Finally, the three last columns of Table 3 show the fixed-effect ordered logit results, when we split the sample by our measure of political institutions based on parliamentary democracy. Column (5) coincides with column (1) in Table 2 and is only reproduced for convenience. As columns (6) and (7) demonstrate, the positive effects on violence of lower wages due to natural disasters and higher aid due to cold-war security-council membership are only present in the countries with non-inclusive political institutions.

All in all, the results in these three tables are remarkably consistent with the theoretical predictions discussed in Section 2.

5 Final Remarks

In this paper, we take a few steps to integrate two different strands of research on political violence, on the determinants of repression and of civil war. We develop a theoretical model to analyze the common roots of both forms of political violence. Under specific conditions on the conflict technology, peace, government repression (one-sided violence) and civil war (two-sided violence) become ordered states depending on a common latent variable. This variable is affected by some important economic shocks: to the value of public goods, aid, wages and resource rents. These effects kick in when political institutions do not provide enough checks and balances or enough protection for those excluded from power.

Another centerpiece of the paper is to bridge the gap between theoretical modeling and econometric testing. Under specific assumptions on observability, our model predictions can be taken to the data by estimating either an ordered logit, or the conditional probability of transition from peace to violence and the transition from non civil war to civil war.

Our empirical identification relies on three sources of, arguably, exogenous variation in the determinants of violence isolated by the theory: wages (via natural disasters), aid (via U.N. security-council membership), and resource rents (via world-market prices of exported commodities). The empirical results are consistent with the theoretical predictions when we think of these variables as driving the latent variable Z in the model. Higher government incomes or lower wages raise the likelihood of government repression, as well as civil war, but only if political institutions have bad checks and balances or bad representation.

The findings in our paper resonate with prior contributions emphasizing the role of institutions, economic development and natural resources in affecting civil conflict, or political violence more generally. However, much work remains to complete our 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. It can also be extended so to make groups different 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. Preliminary investigations in this direction suggest that the impact of heterogeneity on political violence is subtle and less clear-cut than is often claimed in intuitive reasoning.

Our empirical analysis 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 an underlying source of state dependence making the model genuinely dynamic. Such progress could be made by introducing asymmetry between groups. The state variable would then be the group in power, which would make 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.

More generally, it would be interesting to study – theoretically as well as empirically – the two-way links between political violence and economic development. This is a difficult issue, but a possible approach is suggested by Besley and Persson (2009b), who use a simple version of the framework in Besley and Persson (2009c) to study the interactions between political conflict and the building of state capacity, where the development of the state goes hand in hand with development of the economy.

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6 Appendix

6.1 Proof of Proposition 2

Proof. To simplify the notation, the proof leaves out the dependence of γ on parameter vector $\boldsymbol{\xi}$. 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

$$\begin{aligned}\left[2\gamma_O\left(L^O,L^I\right)(1-2\theta)\left[Z-L^I\right]-1\right]L^O &= 0 \text{ for } L^O<X\\ \left[2\gamma_O\left(V,L^I\right)(1-2\theta)\left[Z-L^I\right]-1\right] &\geq 0 \text{ otherwise .}\end{aligned}$$

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, observe that differentiating and using the first-order conditions yields:

$$\begin{bmatrix} -\frac{\gamma_{II}}{2\gamma_O}+2\gamma_I(1-2\theta) & \gamma_O(1-2\theta)-\frac{\gamma_{IO}}{2\gamma_O} \\ \frac{\gamma_{IO}}{\gamma_O}-2\gamma_O(1-2\theta) & \frac{\gamma_{OO}}{\gamma_O}\end{bmatrix}\begin{bmatrix} dL^I \\ dL^O\end{bmatrix}=\begin{matrix}\gamma_I(1-2\theta)\,dZ\\ -2\gamma_O(1-2\theta)dZ\end{matrix}\ . \tag{13}$$

Define $\Omega=\frac{\gamma_{OO}}{\gamma_O}\left[-\frac{\gamma_{II}}{2\gamma_O}+2\gamma_I(1-2\theta)\right]+2\left[\frac{\gamma_{IO}}{2\gamma_O}-\gamma_O(1-2\theta)\right]^2>0$. Solving (13) using Cramer's rule yields:

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

and

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

where we have used both parts of Assumption 1 part d.

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

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

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

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

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

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

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

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

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 1c for all values of θ.

Finally, it is easy to see from the explicit definition that $Z^I(\theta)$ is a decreasing 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 decreasing as well. This concludes the proof of the proposition. ■ ■

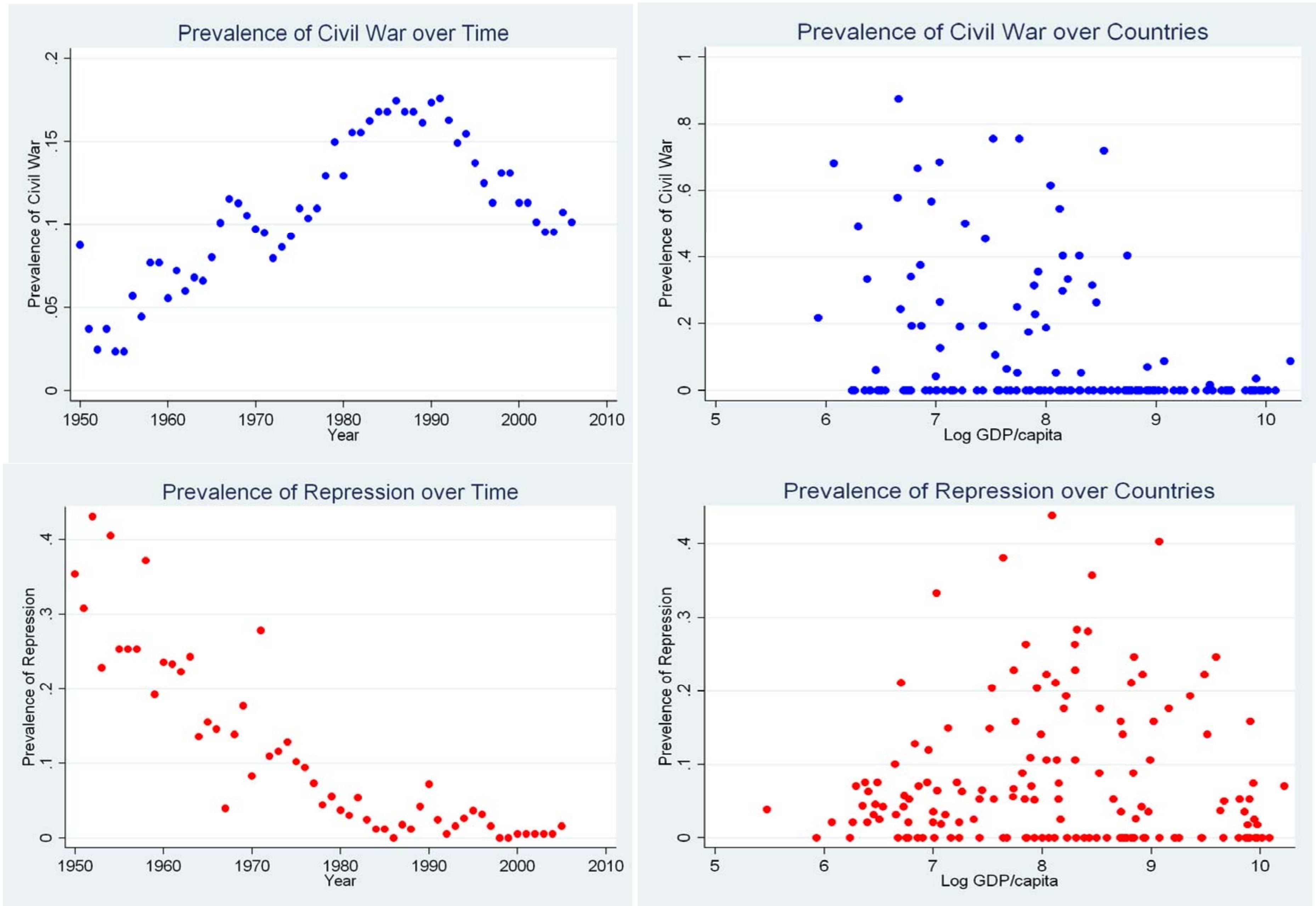

Figure 1 Prevalence of civil war and repression

Table 1 Economic Shocks and Political Violence

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Dependent Variable	Political violence	Political violence	Political violence	Civil war	Civil war	Civil war	Ordered variable	Ordered variable	Ordered Variable
Natural Disaster	0.272*** (0.105)	0.342 (0.496)	0.286** (0.109)	0.349** (0.147)	-1.145 (1.085)	0.401** (0.151)	0.281** (0.106)	0.363 (0.498)	0.296** (0.110)
Sample	All 1950-2005	High executive constraints 1950-2005	Low executive constraints 1950-2005	All 1950-2005	High executive constraints 1950-2005	Low executive constraints 1950-2005	All 1950-2005	High executive constraints 1950-2005	Low executive constraints 1950-2005
Estimation Method	FE Conditional logit	FE Conditional Logit	FE Conditional Logit	FE Conditional Logit	FE Conditional Logit	FE Conditional Logit	FE Ordered Logit	FE Ordered Logit	FE Ordered Logit
Observations	4694	450	4244	2275	114	2161	4382	447	3935
No. Countries	98	8	90	49	2	47	98	8	90

Notes: All specifications include year dummy variables. Standard errors in parentheses: * significant at 10%; ** significant at 5%; *** significant at 1%

Table 2 Economic Shocks and Political Violence: Additional Measures

	(1)	(2)	(3)	(4)	(5)	(6)
Natural Disaster	0.263**	0.602	0.274**			
	(0.107)	(0.536)	(0.111)			
Security council member	-1.048***	37.726	-1.248***	-0.889**	37.388	-1.086***
	(0.399)	(5,298)	(0.416)	(0.405)	(3,618)	(0.420)
Security council member in cold war	1.275***	-39.300	1.560***	1.204***	-38.657	1.457***
	(0.439)	(5,298)	(0.457)	(0.453)	(3,618)	(0.468)
Export price index				0.277**	-54.080***	0.258**
				(0.114)	(19.714)	(0.113)
Sample	All 1950-2005	High executive constraints 1950-2005	Low executive constraints 1950-2005	All 1950-2005	High executive constraints 1950-2005	Low executive constraints 1950-2005
Method of Estimation	FE Ordered Logit	FE Ordered Logit	FE Ordered Logit	FE Ordered Logit	FE Ordered Logit	FE Ordered Logit
Observations	4251	440	3811	4131	368	3763
No. Countries	97	8	89	95	8	87

Notes: The dependent variable is the ordered variable discussed in the text. All specifications include year dummy variables. Standard errors in parentheses: * significant at 10%; ** significant at 5%; *** significant at 1%

Table 3 Alternative Estimation Method and Measurement of Political Institutions

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dependent Variable	Political violence	Civil war	Political violence	Civil war	Ordered variable	Ordered variable	Ordered Variable
Natural Disaster	0.023*	0.029**			0.263**	-0.212	0.311***
	(0.013)	(0.012)			(0.101)	(0.407)	(0.106)
Security council member	–0.066**	–0.051**	–0.051*	–0.043*	–1.048***	1.862	–1.381***
	(0.027)	(0.021)	(0.027)	(0.021)	(0.399)	(1.162)	(0.457)
Security council member in cold war	0.090**	0.034	0.082**	0.024	1.275***	–2.371*	1.668***
	(0.040)	(0.029)	(0.037)	(0.027)	(0.439)	(1.361)	(0.497)
Export price index			0.034**	0.027**			
			(0.016)	(0.012)			
Sample	All 1950-2005	All 1950-2005	All 1960-2005	All 1960-2005	All 1950-2005	Parliam. Democracies 1950-2005	Non-Parliam. Democracies 1950-2005
Estimation Method	FE linear probability	FE linear probability	FE linear probability	FE linear probability	FE Ordered Logit	FE Ordered Logit	FE Ordered Logit
Observations	5880	5880	5609	5609	4251	437	3814
No. Countries	158	158	133	133	97	8	89

Notes: Robust standard errors in parentheses: * significant at 10%; ** significant at 5%; *** significant at 1%. All specifications include year dummy variables.