Holding on? Ethnic divisions, political institutions and the duration of economic declines

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Abstract

We analyze the duration of large economic declines and provide a theory of delayed recovery. We show theoretically that uncertain post-recovery incomes lead to a commitment problem which limits the possibility of cooperation in ethnically heterogeneous countries. Strong constraints on the executive solve this problem by reducing the uncertainty associated with cooperative behavior. We test the model using standard data on linguistic heterogeneity and detailed data on ethnic power configurations. Our findings support the central theoretical prediction: stronger constraints on the political executive shorten economic declines. The effect is large in ethnically heterogeneous countries but virtually non-existent in homogeneous societies. Our main results are robust to a variety of perturbations regarding the estimation method, measures of heterogeneity, measures of institutions and the estimation sample.

Keywords: economic crises, delayed recovery, political economy, ethnic diversity

JEL Classification: E61, O11, O43, J15, H12

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1 Introduction

Why are economic declines in Sub-Saharan Africa and some parts of the globe so persistent? In this paper, we propose a novel answer to this question which incorporates two well-known features of the subcontinent: high ethnic diversity and weak political institutions. We offer a theory of how these two interact during economic declines and test its empirical implications. Our main contribution is to outline a simple mechanism which links political heterogeneity and the powers of the executive to the repeated failure to agree on a policy response to an economic shock, even when the policy is economically effective and socially desirable. We explicitly focus on the process of policy formulation during the decline phase of a slump and consider the realities of countries with weak political institutions. This allows us to show that the uncertainty and distributional consequences created by imperfect constraints on the political executive lead to longer declines in ethnically heterogeneous countries.

Every crisis creates winners and losers. Our model highlights a commitment problem among those that benefit and those that suffer during the recovery process. *Ex ante* uncertainty about post-recovery incomes and a 'winner-take-all' effect caused by weak political institutions can lead to delays in the policy response. Delayed cooperation happens because ethno-political groups with political influence want to limit the risk of being expropriated, or exploited in some other way, by fortifying their own position. While we leave the precise characteristics of the policy response implicit, we assume that these groups are bargaining over some stabilization policy with between-group distributional consequences, such as a nationalization of a particular sector or an investment program.

We derive three major insights from the model. First, delayed cooperation can occur in equilibrium, and weak constraints on the executive act as a political friction in ethnically diverse countries that can lead to social inefficiencies. Second, stronger political institutions can resolve this issue and bring about cooperation early on. Third, all else equal, the commitment problem between winners and losers is worse when the number of groups is large, and the introduction of institutional imperfections has a larger negative effect in more heterogeneous countries. We also outline an additional result

which takes the relative size (strength) of the groups into account to show that political concentration matters for delayed cooperation.

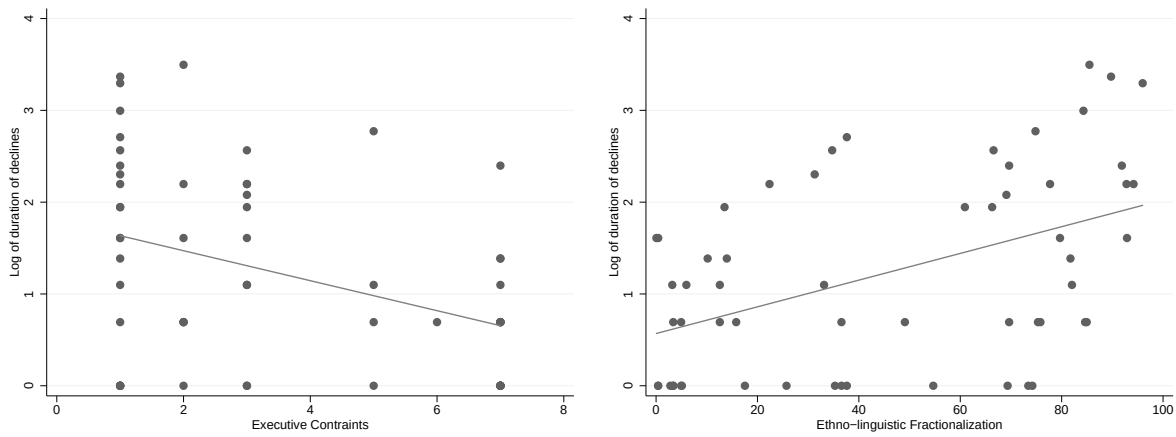
Next, we take the model to the data. We first present a variety of partial correlations consistent with the proposed theory. We examine the central predictions using both standard data on linguistic heterogeneity and a more detailed data set which codes how much access ethnic groups have to the political executive. In line with our theory, we find that the effect of executive constraints on the length of declines is very large in ethnically heterogeneous countries, but muted in ethnically homogeneous countries. This result is robust to many perturbations (e.g. different data sets, different measures and forms of heterogeneity, region and decade dummies, changing the functional form). We also show empirically that greater political concentration shortens declines and, *vice versa*, that a more even distribution of political power across groups increases delay. Possibly counter to intuition, we find that societies in which a larger group dominates or monopolizes the decision-making process experience shorter declines even when executive constraints are weak. In addition, we test a key assumption of the model and show that the number of ethno-political groups represented in the executive not only decreases during the decline phase of a crisis but also during the early years of a recovery.

Our empirical results are not exclusively driven by Sub-Saharan Africa even though the countries on the subcontinent are on average very diverse, institutionally weak and tend to experience the longest declines. Even when we only use within-region variation, our coefficients of interest are statistically significant and substantively large, while the Africa indicator is always significant. Theoretically, we start from the premise that political power lies in the hands of the executive which distributes cabinet seats along ethnic lines in proportion to population shares. This broadly reflects the situation in Sub-Saharan Africa (Francois et al., 2015) but we believe the theory captures how uncertainty and diversity interact during a crisis more generally. Outright expropriation of ethno-political groups is an extreme form of ethnic favoritism, which is often more subtle, but tends to decrease with stronger political institutions (e.g. Franck and Rainer, 2012; Hodler and Raschky, 2014; Burgess et al., 2015). However, such 'winner-take-all' effects,

often linked to African politics, only operate at the margins in our model and can be offset by stronger constraints on the executive. An important policy implication is that well-designed (and enforced) political institutions are key to containing the adversarial element of ethnic diversity and thus play a critical role in many developing countries.

Figure 1: Unconditional correlations with the duration of declines

(a) Executive Constraints (b) Ethno-linguistic Fractionalization



Note(s): The figure plots the log of the duration of economic slumps over executive constraints in panel a) and linguistic fractionalization in panel b). The durations are based on the 58 slumps (in 51 countries) estimated using the approach outlined Bluhm et al. (2014). The duration of the decline phase is the time from the downbreak until the trough. No adjustment has been made for censored observations (unfinished declines). Constraints on the political executive are measured using an index scaled from 1 to 7 (least to most constrained) from the Polity IV data, and ethnic heterogeneity is proxied for by an index of ethno-linguistic fractionalization scaled from 0 to 100 (Desmet et al., 2012).

Our dependent variable is the duration of economic declines. In Sub-Saharan Africa, this duration is five times longer than in Europe and twice as long as the world average. In line with our theory, we specifically focus on the duration of the decline phase of unexpected and large crises, whose beginning is defined by a negative structural break from a previously positive growth trend. Such crises have large welfare consequences and can easily wipe out more than a quarter of GDP per capita in the course of several years (as in Mozambique, 1981–1986, or Zambia, 1968–2001). The methodology behind the econometric identification of these slumps is developed in a recent empirical contribution (Bluhm et al., 2014). That paper deliberately focuses on the duration of the decline phase for three reasons. First, the start of a crisis is often idiosyncratic and not necessarily related to a country's political institutions or level of social cohesion, but its duration depends on socioeconomic groups agreeing on coordinated responses. Second,

the dynamics of recoveries differ a lot from the dynamics of declines (both empirically and theoretically). Third, most of the variation in the overall depth of slumps is due to the duration of the decline segment and not due to the rate of contraction. The motivating observation for this paper is that duration (in years) until a recovery starts increases with greater ethnic divisions and decreases with stronger constraints on the executive. Figure 1 illustrates the unconditional correlation of the (log) duration of declines with executive constraints (-0.39) and ethno-linguistic fractionalization (0.47). Bluhm et al. (2014) are primarily concerned with the econometrics of identifying declines and establishing these stylized facts. The main objective of the current paper is to propose a theory that can explain these observations and to empirically examine our theoretical predictions using detailed data on ethnic groups and their access to political power.

Our work is also motivated by a growing literature which emphasizes that economic growth is often not steady but instead characterized by different growth regimes. For example, it is well known that the correlation of growth rates across decades is low (Easterly et al., 1993). A key finding of the growth episodes literature is that growth accelerations are triggered by a variety of factors but are difficult to sustain (Hausmann et al., 2005; Berg et al., 2012). In developing countries, several years of positive growth can easily be followed by long and deep slumps. Such negative shocks can cancel out previous welfare gains and are often characterized by persistent output loss (Cerra and Saxena, 2008). This volatility starts to play a role in recent institutional theory, such as the 'limited access orders' of North et al. (2009). Little is known, however, about the deeper, more structural factors that are associated with longer (or shorter) slumps.

It is well established that ethnic heterogeneity is a fundamental determinant of economic prosperity. Heterogeneity is typically associated with low growth (Easterly and Levine, 1997), the undersupply of public goods (Alesina et al., 1999), and civil conflict (Fearon and Laitin, 2003; Esteban and Ray, 2011; Esteban et al., 2012). Ethnicity plays a rampant role in Sub-Saharan Africa where political organization is mostly ethnic, but diversity has also been linked to inadequate public good provision in US states (Alesina et al., 1999) or excessive deforestation in Indonesia (Alesina et al., 2014).

Furthermore, high economic inequality among ethnic groups is associated with regional underdevelopment and political inequality (Alesina et al., 2016). Yet the role of ethnic heterogeneity during economic downturns has not been explored.

Heterogeneity is, however, not necessarily a problem and is viewed favorably in many literatures. In developing economies, organizing along ethnic lines may resolve a contracting problem and help to enforce social sanctions within family or kin groups (Bates, 2000). In highly developed economies, the negative effects of ethnic heterogeneity may become muted, as skill complementarities matter more, or political institutions tame the conflict element inherent in diversity (Alesina and Ferrara, 2005). We formally incorporate this latter channel by showing that the negative effects of diversity on cooperative outcomes depend on the strength of political institutions.

Ethnic diversity and weak political institutions often coincide. On the one hand, the adverse effects of ethnic heterogeneity may only be relevant in weakly institutionalized societies where political leaders often use (or abuse) ethnic and other divisions in their favor (Eifert et al., 2010). On the other hand, diversity affects the (endogenous) choice of institutions governing the executive power of such leaders (Aghion et al., 2004). There is some empirical evidence consistent with the view that ethnicity and political institutions interact.[1] However, the precise mechanisms behind how these two jointly determine the length of crises have not been investigated and may explain *substantial* parts of the robust negative correlation between ethnicity and growth. While plenty of anecdotal evidence exists, we are only aware of a paper by Rodrik (1999) which explicitly considers ethnicity and negative growth empirically (and more formally in the working paper version).

The stylized facts motivating this paper cannot be explained with established theoretical frameworks. Ethnic groups could be engaged in "wars of attrition" over the burden of reform, so that groups are trying to shift the costs of, say, a debt consolidation onto competing groups (Alesina and Drazen, 1991).[2] Alternatively, a socially optimal

[1] Collier (2000), for example, argues that ethnicity plays no role in democracies but reduces growth in autocracies and provides evidence along these lines. Easterly (2001) empirically investigates an interaction effect between institutions and ethnicity in determining growth and conflict.

[2] In these models, groups learn about the capacity of their opponents to bear the costs of waiting as time passes and stabilization occurs only once one of the groups concedes. Drazen and Grilli (1993) use this set-up to show that crises can be welfare improving by reducing delay. Spolaore (2004) examines

reform may not be undertaken at all because it is *ex ante* not known to which (ethnic or other political) groups the benefits will accrue (Fernandez and Rodrik, 1991). Such a model can also generate delay and an endogenous economic deterioration (e.g. Labán and Sturzenegger, 1994). Both approaches have two key elements in common: 1) *uncertainty* about the expected outcomes, and 2) an *ex ante commitment problem* between (*ex post*) beneficiaries and losers of the reform. While instructive, these models are not ideally suited for the setting we consider here, where a crisis is immediately obvious and the pre-crisis political power of each key player is often well known. Furthermore, these papers do not explicitly focus on ethnic diversity and presuppose the existence of strong political institutions. In addition, wars of attrition should generally predict that greater constraints on the executive prolong the time to stabilization, as such constraints make it easier for weak players to hold out. This is the opposite prediction with respect to institutional strength than the relationship we highlight theoretically and find empirically.

Another popular framework through which related questions have been analyzed is the veto player literature in political science. Contributions based on this framework generally find that policy stability is greater the more numerous the players in the political system that are required to agree (e.g. Tsebelis, 2002). Veto player arguments have been used to explain why governments may not reform during an economic shock (e.g. Cox and McCubbins, 1997). Within this literature, Hicken et al. (2005) come closest to our paper with their emphasis on political institutions and economic performance. However, they conclude that greater checks on the executive do not aid recovery, which stands in sharp contrast to our empirical findings.

The level of ethnic diversity is endogenous in the (very) long run. Heterogeneity is related to migratory distance from Africa (Ashraf and Galor, 2013), the duration of settlements and the history of the state (Ahlerup and Olsson, 2012), and variation in terrain and land endowments (Michalopoulos, 2012). At the micro-level, people may choose their group affiliation and switch groups depending on how discernible the individual features are which identify group membership (Caselli and Coleman, 2013).

the impact of different government systems on the expected time until a stabilization occurs.

However, we do not expect ethnic compositions to change fundamentally in the short run (especially in the post-colonial period). That said, ethnicity is not always the most prominent political fault line in a society and the degree of access to political power of a particular group varies over time (Posner, 2004). Early empirical studies of the effects of ethnic heterogeneity (e.g. Easterly and Levine, 1997) use data Soviet ethnographers published in 1964, incorporate possibly irrelevant cleavages and do not account for differences in political power. Several later studies use up-to-date data on *linguistic* fragmentation (e.g. Fearon, 2003; Desmet et al., 2012) but still remain confined to the cross-section and disregard political power. Wimmer et al. (2009) present a new data set which explicitly aims to remedy this situation. The *Ethnic Power Relations* (EPR) data codes the degree of access to executive power by different groups, focuses on politically relevant groups, and employs a more flexible notion of political division capturing the main fault line in a particular country (such as ethnicity, language, race or religion).[3] In our empirical part, we use the latest EPR data to unpack the model's predictions and contrast our findings to other established measures of diversity.

The rest of the paper is organized as follows. In Section 2, we outline our model of how ethnic fractionalization and weak constraints on the executive can lead to delayed cooperation. In Section 3, we discuss the data, the empirical strategy and the main empirical results. Section 4 concludes.

2 Theory

We model group interactions during a slump as a cooperation game where groups decide on whether to formulate a policy response to a crisis that will initiate a recovery. We first focus on the symmetric two group case and then extend the model to allow for unequal sizes and multiple groups.

[3] Francois et al. (2015) extend this approach further and provide data on the ethnic composition of the ministerial level in 15 African countries. If this data were available for more countries across the globe, then it would be an ideal supplement to the EPR data used here.

2.1 Basic setup

We consider a population normalized to unity and split into J equal-sized (ethnic) groups. These J groups constitute the players of the game. Time is discrete and there is an infinite number of periods, indexed by $t = 1, 2, \ldots, \infty$. The per-period discount rate is δ. With slight abuse of notation, groups are indexed by $j = 1, 2, \ldots, J$, where $J = 2$ for the baseline model considered in this subsection. Each group acts as a single agent and we do not analyze internal coordination issues among members of the same group.

Preferences. Group j receives a net income of y_j in period zero. Total initial income in the economy is normalized to unity ($\sum_j y_j = 1$). Utility in each period is $g(y_j)$, where $g(\cdot)$ is increasing in y_j, concave and identical for all groups.

Slumps: decline and recovery. When a slump occurs, output declines by a fixed amount (Δ) in the first period. The income shock affects both groups proportionally and output remains at that level until both groups cooperate. Total income is now $1 - \Delta$ as long as the slumps lasts. Once a decision to cooperate has been reached, we assume that the economy recovers within one period. Groups decide to cooperate or not based on their expected future returns to cooperation.

We leave open the exact nature of the actions that can be taken to facilitate recovery. One example would be the implementation of a stimulus package in an economy well below potential output, possibly involving conditional loans from international financial institutions. Another possibility would be a bailout or nationalization of a banking sector at the verge of collapse or a bailout (nationalization) of a commodity sector accounting for a non-trivial share of the economy. Note that all of these policies are likely to have implications for the economic and political power of the affected groups.

Slumps: uncertainty. We assume that groups are uncertain about their post-recovery incomes – their relative economic standings and political power may change after the slump is over. In the baseline model, $y_1 = y_2 = 1/2$. Each group experiences a random shock to its income, where the probability that a group falls below a "threshold of safety"

and is expropriated by the other group is given by a process p_t (which is explained in greater detail below). The setting is symmetric in the sense that p_t also gives the probability of the second group falling below the threshold. For the first group, the shock has support $\nu_1 \in [-y_1, 1 - y_1]$. Let w_1 denote actual income after the shock, so that y_1 is now a counterfactual; similarly for y_2, w_2. This implies that a slump will hit the groups unequally after recovery, but *ex ante* neither group expects to be hit harder.

Political institutions. We interpret executive constraints as limits on how much one can group gain or lose relative to the other through expropriation, as is common practice in the literature (Besley and Persson, 2011a,b). The intuition is as follows. If a particular group has been sufficiently weakened by the slump, the now stronger group may be in a position to expropriate part or all of the weaker group's income and exclude it from the political process. If the executive is unconstrained there are no checks on this type of predatory behavior. A complete constraint on how much one group can extract means that no expropriation can occur. A partial constraint implies that expropriation occurs only when one group becomes too weak. The dominant ethnic group controls the executive and shares the spoils from expropriating the weaker group with its members.

To be more precise, we model political institutions by including thresholds in the random shock. This introduces a second source of uncertainty: boundary outcomes (expropriation) are realized only beyond these thresholds. Let the parameter $c \in [0, 1/2]$ represent the weakness of executive constraints, and let the set $\mathcal{A} = [c, 1 - c]$ be the political "safe zone" in which there is no expropriation. Once a group falls below c, its income is expropriated (pushed to zero) and the other group gains the remainder. Thus, $1 - 2c$ can be interpreted as the ability of one group to commit to not expropriating the other group; alternatively, $c = 1/2$ can be thought of as the total lack of constraints.

To fix ideas, we interpret the 'winner-take-all' event as political extinction of the weaker group, though it can be understood in a variety of ways. In non-democratic politics, the threshold mechanism symbolizes the potential of some ethnic groups to exclude other groups from the political process and capture the rents of those that have been excessively weakened by the slump. Alternatively, it may even represent

physical extinction due to ethnic conflict. In democratic politics, assuming that ethnic or other identity groups are represented by parties reflecting their interests, it captures the existence of thresholds that allow minorities to participate in government (e.g. the filibuster rule used in the U.S. Senate as well as several state legislatures, or, perhaps more fittingly in terms of identity politics, the 10% electoral threshold used in Turkey's general elections[4]).

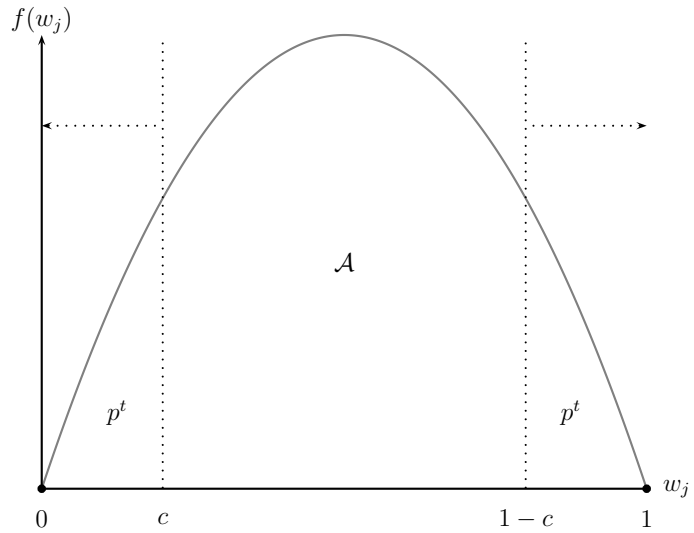
Delay. We assume that groups are able to fortify their position through non-cooperation. This implies that a group can (in part) counterbalance the uncertainty introduced by weak institutions through not cooperating, and thus potentially avoid falling below the threshold. The fortification of positions can be interpreted in multiple ways, with the appropriate interpretation depending on the context. In some countries, it can mean literal fortification, with political leaders mobilizing loyal members from an ethnically or regionally defined group for a show of force to the country's capital, or elsewhere to protect valuable resources. In societies less prone to violence and civil conflict, it can be interpreted as the building of political alliances, or the moving of production activities and resources to safe places (possibly outside the country). Such actions limit the risk associated with a change in economic and political influence, but, most importantly, *they take time*.

In terms of the model, delay limits how likely it is that a particular group will be expropriated. The parameter x is a measure of how much a group can reduce the risk of expropriation by holding out in each period. We assume that the probability of landing on either side outside the safe zone follows a linear process, so that $p^t = c - (t-1)x$ at each t when the groups can chose to cooperate or delay. Furthermore, we assume that expected utility conditional upon being in the safe zone is independent of p^t.

Figure 2 gives an example of a distribution of w_j and illustrates the relevant regions.

[4]This rule often kept Kurdish minority parties, but also other established parties, out of parliament. In the 2002 national elections, 46.33% of all votes were cast for parties below the 10% threshold and hence not represented in parliament.

Figure 2: Threshold effects as constraints on the executive



Note(s): The figure provides an example of a probability distribution of after-shock incomes $f(w_j)$ in the two group case. The weakness of executive constraints is given by the thresholds c and $1-c$. The safe-zone is the area denoted by $\mathcal{A}$. The probability of falling below or above the threshold at time t is denote p_t. A group which falls below the threshold c on the left will be expropriated so that $w = 0$ and the other (stronger) group ends up with $w = 1$ in return.

Timing. The following timing summarizes the structure of the game. At $t = 0$, the economy is in its initial state. Output $\sum_j y_j = 1$ is produced and shared equally.

1. At $t = 1$, the slump occurs, and incomes decline to $(1-\Delta)y_j$. Both groups simultaneously choose to cooperate C or delay D.

2. For all $t > 1$, incomes remain at $(1-\Delta)y_j$ if both groups did not cooperate in the previous period. They once again simultaneously choose whether to cooperate C or delay D. If, instead, there was cooperation in the previous period, incomes recover within one period, but are subject to a random shock and groups can land outside the political safe zone with twice the probability p^t. After a recovery, each group receives the same payoff as in the first post-recovery period forever.

The present discounted value of the lifetime utility for each group is

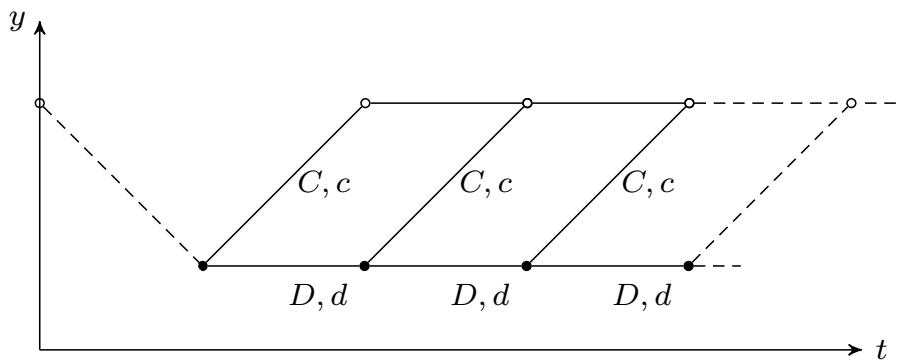
$$v_j = \sum_{\tau=1}^{\infty} \delta^{\tau-1} \mathbb{E} g(\cdot) \tag{1}$$

where $g(\cdot)$ is $g((1-\Delta)y_j)$ if the recovery has not yet occurred and $(1-2c)\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] + c(g(0) + g(1))$ otherwise. The discounted utility has two components: 1) if the

economy has not recovered, groups are on a delay path, and 2) once the slump is over, they remain on a post-recovery path.

Figure 3 sketches how the economy evolves over time given different choices and presents a stylized view of the process we envision. Note that the action pair (D, d) has the same implication as (C, d) and (D, c); that is, cooperation of both groups is required for a recovery to occur.

Figure 3: A sketch of decisions and timing



Note(s): The figure shows a pseudo-game tree where the time path of each groups income (y) is displayed along side the solid choice nodes. When both groups cooperate, a recovery takes place and the same income is received forever on the post-recovery path. The level of this post-recovery income can differ for each group. When one or both groups choose delay, then the recession utility is received for another period, after which both groups will choose again.

The game has a symmetric structure. At each choice node (solid nodes), the comparison between any two adjacent periods always looks alike. The utility from cooperating in a particular period t when the other group cooperates in period t is

$$v_j^t(C, c) = \frac{1}{1-\delta}\left\{(1-2p^t)\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] + p^t(g(0)+g(1))\right\} \tag{2}$$

and the utility from cooperating in the next period when the other group cooperates in period t is

$$v_j^t(D, c) = g((1-\Delta)y_j) + \frac{\delta}{1-\delta}\left\{(1-2p^{t+1})\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] + p^{t+1}(g(0)+g(1))\right\}. \tag{3}$$

It is useful to establish the social optimum before we characterize the non-cooperative equilibrium. Our first comment summarizes two key aspects of the planner's solution.

Comment 0. *i) The utilitarian welfare-maximizing outcome involves no delay.*

To see this, note that due to the concavity of the utility function the sum of the group's utilities is maximized when their share is equal. At equal shares, the total welfare from any non-delay path dominates any delay path.

ii) Any outcome with delay is Pareto dominated by some outcome without delay.

To see why this is the case, take any path with delay, give the groups the same shares in every period, but let the recovery happen immediately. In this case, all groups receive more in the period before the recovery than they did with delay, and the same in every period after the recovery.

The intuition behind this comment is straightforward. Given that there are two groups in the economy, a social planner would give both the same shares and avoid delay; only then is their combined utility maximized. Even if these two groups have unequal shares, an immediate recovery is beneficial to both. The social planner is unconstrained, in the sense that the solution involves no uncertainty towards the post-recovery utilities or political boundary effects. This benchmark is particularly interesting when contrasted to the non-cooperative equilibrium of the game, where groups face a trade-off between immediately recovering and falling below the threshold c, or recovering later and reducing future uncertainty.

By comparing the utilities from cooperating in the first period and in the second period it is relatively straightforward to show that delay can occur in equilibrium. Our first result establishes this.

Proposition 1. *There exist parameter values, such that all equilibria involve delay.*

Proof. See Appendix. ■

The proof to the proposition shows that all components that make the immediate

cooperation scenario less attractive are conducive to delay. The key issue rendering the cooperative equilibrium inaccessible is the *ex ante* commitment problem among potential winners and losers. Hence, worse institutions, or less ability to commit to not expropriate the loser (larger c), larger gains from holding out (larger x) and a larger value placed on the future (higher δ) make immediate cooperation less likely. Conversely, a larger shock (Δ) makes cooperation more attractive since a (potentially sizable) one period loss is avoided. The concavity of $g(\cdot)$ matters in the sense that it implicitly captures how averse groups are to negative events (falling below c) or how much they value expropriating other groups (landing above c).

Note that the proposition is formally true only in a weak sense; it does not rule out that equilibria with immediate recovery could exist for some parameter values.[5] Rather, the result should be viewed in light of Comment 0. What Proposition 1 establishes is that for some parameter values *all* equilibria are inefficient and welfare-suboptimal.

While still in the two-group case, we can already highlight an interesting comparison to the homogeneous (one group) case.

Comment 1. *Without heterogeneity, there always exists an equilibrium with immediate recovery.*

Note that if the groups were to pool their resources as one, then all the elements inducing delay – except pure miscoordination – are absent. In other words, we need antagonistic political (ethnic) groups for the proposed mechanism to work, i.e. for the model developed here to provide a theory of why there is delay. A more careful analysis of group asymmetries and multiple groups follows in the model extensions.

To better understand when we are likely to see delay, we now characterize the subgame perfect equilibrium with (the earliest possible) recovery, if such an equilibrium exists. Given the symmetric structure of the game an interior solution exits and the optimal

[5]There are many "coordination failure" equilibria where neither group cooperates simply because they believe the other group will not. Such equilibria always exist, including an equilibrium with infinite delay. Our analysis, however, is focused on the more interesting scenarios (equilibria) where delay does not happen *only* as a result of this type coordination failure.

time to recovery can be derived using equations eq. (2) and eq. (3). Our second result summarizes a central insight of the model.

Proposition 2. *Stronger constraints on the executive shorten the time to recovery.*

Proof. See Appendix. ■

The proof shows that the optimal time to recovery is

$$t^* = \frac{g((1-\Delta)y_j) - \mathbb{E}[g(w_j)|w_j \in \mathcal{A}]}{x\{2\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] - (g(0)+g(1))\}} + \frac{c}{x} + \frac{1}{1-\delta} \tag{4}$$

where the key comparative static result is $\partial t^*/\partial c > 0$.

This proposition says that if institutions are imperfect ($c > 0$), delay is going to be longer than if the groups are able to perfectly commit to not expropriating the losers.[6] In fact, the weaker the constraints on the executive (larger c), the longer is the expected time to cooperation. Intuitively, either group will find it optimal to delay until a point is reached when the benefit of holding out for an additional period is equal to the benefit of cooperating in this period, where the former may be the period in which all uncertainty regarding the political threshold is resolved. At this point, or the next discrete period, it is optimal to cooperate. Where exactly this point in time occurs depends on the trade off between recovering and potentially falling outside the political safe zone, or recovering later and reducing the remaining uncertainty.

For the remainder, we do not explicitly derive this equilibrium solution. Instead, we focus on the case where all uncertainty is resolved in the next period and compare different scenarios (e.g. perfect and imperfect institutions). We outline such an argument in the next comment.

[6] Again, this is only holds if we rule out equilibria involving immediate cooperation or infinite delay.

Comment 2. *The existence of imperfect (weak) institutions makes delay more likely.*[7]

If institutions are perfect ($c = 0$), we have

$$\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] \geq g((1-\Delta)y_j) \tag{5}$$

and if institutions are imperfect (and $p^t = c$ for all t), we have

$$(1-2c)\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] + c(g(0) + g(1)) \geq g((1-\Delta)y_j). \tag{6}$$

Inequality (6) is harder to satisfy than inequality (5) under the concavity assumptions imposed on the utility function. Note that this is entirely due to the presence of weak institutions ($c > 0$).

Discussion of the model. The baseline model focuses on several key aspects of the political economy of declines. First, we have modeled group interactions during crises under uncertain post-recovery incomes in a way that highlights that groups are not able to commit to compensating the losers. There are no enforceable contracts where the winners return the (additional) post-recovery gains, which is precisely the role played by strong constraints on the executive. Second, outcomes with delay can occur in equilibrium, and they do *not* coincide with the social optimum or with efficiency. Weak institutions act as a political friction creating potentially large economic inefficiencies. Third, heterogeneity matters and political groups are assumed to be willing to cooperate once it is optimal to do so. Entrenched distrust would only increase delay.

The model in this paper is developed specifically for understanding slumps (and their ending). In principle, one could imagine using a similar theoretical framework to predict the onset of other changes in the pace of growth, such as accelerations or extended

[7]Strictly speaking, a probabilistic statement (delay becomes "more likely") should not be used in this comment, as, for any given set of parameters, there either exists an equilibrium with immediate recovery or not. However, we follow Acemoglu and Robinson (2006) and say that a change in the setup of the model makes a particular outcome "more likely" if it becomes an equilibrium outcome for a larger parameter set.

periods of modest but successful development. However, growth accelerations, almost by definition, usually start from periods of unremarkable growth. They are unlikely to be preceded by the sense of urgency and extreme pressure to coordinate that provides the backdrop for the model we develop here. We also do not expect growth spurts to coincide with an elevated risk of expropriation, another key feature of our model, which we later document in the empirical section. With this in mind, we find it prudent to limit our claims of generality and present the model as a theory of economic declines.

We abstract from several features that would be potentially important in a paper with a different focus. For example, we assume the decline does not deepen after the first time period in a slump, and we assume that recovery, once it takes place, is immediate. Assuming an indefinitely continuing decline phase and non-immediate recovery would lead to a more realistic setup of the model. However, while these assumptions would add pressure to agree early on in a manner that might better reflect how slumps actually occur, the focus of our theory part is to understand the qualitative impact of heterogeneity and political institutions; we do not aim for quantitative predictions of the exact time to agreement. For similar reasons, we do not model the precise nature of the policy response, differentiate between democratic and autocratic regimes, or examine the impact of particular political constitutions (presidential or parliamentary). The exact form of the boundary events is also left open and could, for example, also represent the exclusion from public goods. We also do not differentiate between political and economic power. Again, such specificities are not essential to the main argument. Leaving them out does, however, imply that our paper might be a better description of some countries than others. The model is likely to be most relevant for understanding declines in Africa and other countries where political divisions are often ethnic and executive power is shared (e.g. see Francois et al., 2015).

As a final comment on the baseline model, note that the mechanism we propose is different than those suggested in the policy reform literature, which has previously focused on shifting the burden of reform (Alesina and Drazen, 1991) and status-quo bias (Fernandez and Rodrik, 1991). While (some of) the papers in this literature highlight

the importance of *ex ante* uncertainty (either about the costs or benefits of reform), their core focus is not on the role of political institutions in general or executive constraints in particular. The empirical content also differs substantially from ours. For example, Drazen and Grilli (1993) stress that crises help stabilizations and Spolaore (2004) shows that political systems with a strong government (less constrained executive) reform more quickly, whereas we propose that crises coinciding with an unconstrained executive are at the heart of the problem.

2.2 Extensions: asymmetric and multigroup settings

We now briefly sketch two extensions. To extend the model to the asymmetric and J-group cases, we make the following simplifying assumptions. First, we restrict attention to the uncertainty associated with falling below the political threshold. Specifically, we assume that if a group falls within the political safe zone its share of total economic activity will be equal to its pre-recovery share. Second, we use a piecewise linear utility function, in particular:

$$g(y_j) = \begin{cases} y_j & \text{for} \quad y_j > 0 \\ z & \text{otherwise} \end{cases} \tag{7}$$

where $z < 0$. Furthermore, for the case when there are more than two groups, we assume that at most one group can fall outside the political safe zone. We now work with a more general (continuous) probability function, where we only assume that $dp_t(y_j)/dy_j < 0$ for any given level of the constraint c. Finally, for simplicity, our comparative statics will be done for the case where all uncertainty is resolved after one period of delay.

How do changes in political concentration affect the political equilibrium? Intuition may suggest that smaller groups are more afraid of falling out of the political safe zone, implying that greater asymmetry between groups increases the likelihood of delay. However, our theoretical result suggests that the effect of changes in political concentration can go either way. Several things change in the two-group case if the share of an initially weaker group moves closer to an equal allocation, so that the size of the

previously more powerful group decreases in return. On the one hand, the emboldened group faces a lower probability of being expropriated. In addition, the group also has to forgo more utility in the delay scenario. Both work in favor of cooperation. On the other hand, the group now has more to lose if it gets expropriated and is thus less likely to cooperate. Without imposing further restrictions, the overall direction of the effect is undetermined and depends on the parameter values. We consider this an empirical issue and return to it in the next section. The following result gives the condition that has to hold for greater symmetry to lead to more delay.

Proposition 3. *A decrease in (political) concentration makes delay more likely, if the following condition holds*

$$\Delta + \frac{1}{1-\delta}\left\{\frac{dp_1(y_j)}{dy_j}(z-y_j) - p_1(y_j)\right\} < 0. \tag{8}$$

Proof. See Appendix. ■

Using this condition, we can summarize the circumstances that determine the direction of this effect.

Comment 3. *A decrease in concentration is more likely to work in favor of delay, if the shock is smaller, the future is less heavily discounted, the negative consequence of falling outside the political safe zone is greater and the probability of that event is not very responsive to the weaker group's share.*

Up until this point, we assumed that there are only two groups deciding on whether to cooperate or not. The final proposition relaxes this constraint and highlights two key insights of the model with respect to group heterogeneity (assuming symmetric groups).

Proposition 4. *i) An increase in the number of groups makes delay more likely.*
ii) Introducing imperfect (weak) institutions is more likely to lead to delay if the number

of groups is larger.

Proof. See Appendix. ■

Contrary to the more equivocal result in Proposition 3, a larger number of groups decreases the likelihood of cooperation. The proof of part *i)* shows that the condition for immediate cooperation (when all uncertainty is resolved in the next period) boils down to an inequality that decreases in J. The intuition behind this proposition is simple. As the number of groups increases, every group becomes poorer and thus more vulnerable during a slump. Simplifying the model helps to show that this is driven by the uncertainty arising from the lack of executive constraints (which we now implicitly define through $p_t(y_j)$). Part *ii)* then takes the multi-group extension back to the motivating question behind Proposition 1; that is, we again explore why (when) Pareto-inefficient delays can occur, but now with a focus on the interaction between institutions and heterogeneity. It extends the first part of the proposition, which shows that more groups make cooperation less likely, and adds that an introduction of imperfect institutions will be particularly problematic when the number of groups is large.

So far we did not explicitly consider political power or political relevance. Instead we assumed that all groups start from inside the political safe zone, matter equally for the decision to cooperate, and may only fall into political irrelevance as a consequence of the slump. Keeping the decision mechanism fixed, we now reflect on what this implies for different power (group) configurations. We do so with an eye to the concepts that we can empirically capture in the next section. With this in mind, we summarize the role of political relevance in the last comment as follows.

Comment 4. *More politically relevant groups make delay more likely, while politically irrelevant groups do not matter. Conversely, this implies that if a group dominates or monopolizes the decision making process, delay becomes less likely.*

Here, political relevance refers to being a party to the negotiations and thus being included

in this model. Note that this separates the issue of political relevance (and the number of groups) from the strength of constraints on the executive. For example, there may be a dominant group which is unconstrained and thus poses a threat for smaller groups, or there may be a dominant group whose hands are tied so that it cannot expropriate smaller groups. Since these cases are distinct, we should still observe an independent effect for both the number of groups represented in the executive (or different qualitative assessment of their access to executive power), and constraints on the executive *per se*.

As a final point before we move on to the empirical part, we want to highlight that any (positive) effect of executive constraints requires ethnic heterogeneity in our theoretical framework. This follows already from the baseline model. Furthermore, the second part of Proposition 4 shows that, when starting from a welfare-maximizing benchmark, an institutional imperfection that introduces the possibility of delay is more likely to matter in heterogeneous societies. Empirically, we therefore expect that an interaction term between a measure of ethnic heterogeneity and an index of the strength of political institutions should have a *negative sign* (i.e., reduce the duration of the decline phase).

3 Empirical Strategy and Discussion

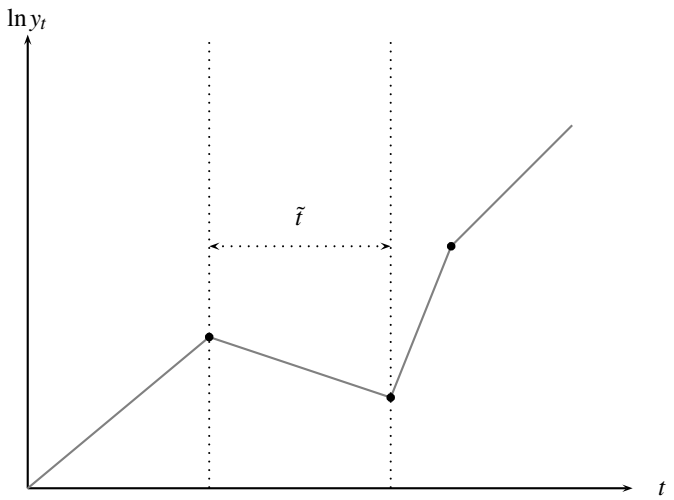
Decline spells. We characterize slumps by an abrupt negative departure from a previously positive growth regime that coincides with two successive trends breaks (usually separating a recovery regime from a post-slump regime). We then calculate the time it takes from the start of the first break until the empirical trough. Hence, our dependent variable is the duration of the decline segment during deep economic slumps. Figure 4 sketches the process we had in mind when designing this algorithm.

Our "to the bottom" definition of slumps differs from the existing literature which typically focuses on successive years of negative GDP growth, the duration from the business cycle peak to the trough, or the duration of entire recession (until full recovery). However, this definition is perfectly in line with our theoretical set-up. It captures our argument that the length of the decline segment depends on the political system's ability

to react to a crisis and accounts for the fact the dynamics of downturns and recoveries are often very different.

Structural break methods are an established way of identifying turning points towards negative or positive growth regimes (e.g. Rodrik, 1999; Hausmann et al., 2005; Jones and Olken, 2008; Berg et al., 2012; Papell and Prodan, 2014). They incorporate a notion of "pronounced" and "unexpected" declines in a univariate time-series sense, and allow us to statistically discriminate among multiple plausible starting points. However, since the identification of the duration of negative growth spells is not trivial and beyond the scope of this paper, we only briefly summarize the method here. More details can be found in the appendix of this paper and are discussed at length in Bluhm et al. (2014).

Figure 4: A stylized decline spell



Note(s): The figure shows a stylized example of the logarithmic time-series of GDP per capita before and after a major slump. The duration of the decline phase, denoted $\tilde{t}$, is the time (in years) between the first break and the through. The second break after the trough denotes another shift in the growth trend typically occurring after a recovery.

The procedure involves several steps. First, we fit a restricted partial structural change model with two breakpoints to each GDP per capita series. We impose sign restrictions on the model parameters, so that we only find major economic slumps. Second, we estimate candidates for the endogenous breakpoints and conduct a bootstrap Monte Carlo test of their significance. Third, we keep only breaks that are significant at the 10%-level and run the procedure again on the remaining data (before the first and after the second break) until all breaks have been found or the sample gets too small. Fourth, for each slump, we identify the empirical trough (the lowest point in the series after the beginning

of the slump) and then compute the duration of the decline segment (denoted $\tilde{t}$). The spell is censored if pre-slump GDP per capita has not been recovered by the end of the sample, since we cannot rule out the possibility that the true trough occurs in the future.

Applying this algorithm to the *Penn World Table 7.0* yields 58 slumps in 51 countries from 1950 to 2008.[8] Table A-1 in the appendix provides a full list. Most slumps occur in the 1970s, 1980s and the early 1990s. The first observed slump begins in 1953 and the last observed slump begins in 1997. The method identifies many well-known slumps in both developing and developed countries, such as the Mexican debt crisis of the early 1980s (followed by the Tequila crisis in 1994) and the Finnish banking crisis of the early 1990s, but also deep and long-lasting declines in a number of African countries (e.g. Mozambique, 1981–1986, or Togo, 1979–2008). For the robustness checks, we also use a more lenient significance threshold of 20%, which results in a larger sample of 83 slumps in 70 countries and also includes many well-known episodes.

Table 1: Summary Statistics of Slumps

	Africa	Americas	Asia	Europe	Oceania	World
Countries with slumps	14	11	15	9	2	51
Number of slumps	14	16	16	9	3	58
Total years in decline	178	78	60	23	9	348
Duration of decline:						
- Min	1	1	1	1	2	1
- Median	16	2	2	1	3	3
- Mean	12.71	4.88	3.75	2.56	3.00	6.00
- Max	33	15	13	9	4	33
Incidence Rate	0.04	0.19	0.22	0.39	0.33	0.14

Note(s): The table shows summary statistics of the duration of the decline phase of slumps. A few countries (e.g. Chile) have repeated slumps which generates the discrepancy between the reported number of countries and number slumps. The incidence rate is defined as the number of exits from the decline period over the total years in decline.

Table 1 provides some summary statistics. The basic correlations are as expected. Poorer countries have longer and deeper declines than richer countries; countries in Africa have the longest and deepest spells. OECD countries do experience their fair share of volatility (12 slumps) but they tend to be shallow and short spells. While the distribution of slumps is relatively even across the different regions, their depth and average duration

[8]Note that we exclude countries with less than one million inhabitants and less than 20 years of data.

varies greatly between developed and developing countries. We only observe a small number of repeated spells, three of which occur in Chile (starting in 1953, 1974, 1981). Ten slumps are unfinished; that is, GDP per capita has not recovered to the pre-slump level by the end of the period under investigation. Their trough is estimated to occur at the lowest observed value of GDP per capita and the spell is censored.

Measuring institutions. Our core measure of political institutions is the variable *Executive Constraints* from the Polity IV data set.[9] The variable directly measures the degree of institutionalized constraints placed on the political executive. It is coded unity when there is "unlimited executive authority" and seven when there is "executive parity or subordination"; intermediate values represent some constraints. We believe that this variable corresponds well with the parameter c in our model. The Polity IV project has information on executive constraints annually from 1800 (or the year of independence) until 2010. We do not use this wealth of time variation, since political institutions endogenously respond to the slump (Bluhm et al., 2014). We only rely on the degree of executive constraints in the last year *before* the slump and denote this variable $XCONST_0$. This rules out the possibility of feedback from the duration of declines to our measure of the risk of expropriation.

Measuring heterogeneity. We rely on two data sources to capture very different aspects of ethnic heterogeneity. The first source is a set of measures computed by Desmet et al. (2012) on the basis of the *Ethnologue* data. This data does not measure ethnicity directly but captures *linguistic* diversity. Fearon (2003) shows that linguistic (cultural) diversity coincides well with ethnic heterogeneity in some regions, notably Sub-Saharan Africa, but not so well in others. Together with the *Atlas Narodov Mira* data gathered by Soviet ethnographers in the 1960s, it is a standard source for data on ethnic heterogeneity and considerably more up-to-date than the former. Desmet et al.

[9] An alternative measure of political constraints is the *POLCON* index proposed by Henisz (2000). This index is not our preferred measure for three reasons. First, the measure is derived from a veto-player model, while we propose a different theoretical approach. Second, it focuses on the number of parties in the legislature, not structural features of the executive. Third, it explicitly includes legislative fractionalization whereas we emphasize ethnic fractionalization of the executive. Additional results using this measure can be found in Table A-6.

(2012) compute linguistic diversity at different levels of the language tree to capture the historical depth of ethnic divisions. We only make use of the most disaggregate level, since they also show that current divisions are correlated with economic growth more strongly than historical cleavages. The second data source is the *Ethnic Power Relations* (EPR) data presented in Wimmer et al. (2009). The EPR data has several advantages over other measures of linguistic or ethnic diversity, particularly for our application. It provides time series information on the degree of access to *executive* power of ethno-political groups from 1946 to 2010. Contrary to the *Ethnologue* data, it is not restricted to linguistic cleavages existing today. Instead, expert coders identify the most relevant division which may be ethnic, linguistic, racial or religious depending on the country and time period. The data contains information on the power status of each group, so that it allows us to focus on politically relevant groups; that is, groups with some form of representation in the presidency, cabinet, or other senior posts in the administration or army.

Our primary measure of heterogeneity is the commonly used index of ethno-linguistic fractionalization (e.g. Easterly and Levine, 1997). It is defined as

$$ELF_i \equiv 1 - \sum_{j=1}^{J_i} \left(\frac{n_{ij}}{N_i}\right)^2 = \sum_{j=1}^{J_i} \frac{n_{ij}}{N_i}\left(1 - \frac{n_{ij}}{N_i}\right) \tag{9}$$

where n_{ij}/N_i is the population share of group j in country i ($j = 1, 2, \ldots, J_i$, n_{ij} is the number of people in group j, and N_i the size of the population in country i). We employ two versions of this index: one computed by Desmet et al. (2012) and one computed using all group represented in the EPR data anchored to its pre-slump value (denoted ELF_0). We scale all heterogeneity indices by 100 to give changes on the right hand side a percentage point interpretation.

Another important dimension of diversity is the degree of polarization of a society. The literature on ethnic conflict often stresses that fractionalization and polarization have very different effects (e.g. see Esteban and Ray, 2011). We capture polarization with an

Table 2: Definitions of Variables

Symbol	Description	Source and Notes
	Dependent Variable	
$\tilde{t}$	Duration of decline segment	From Bluhm et al. (2014) computed using structural break model with a significance level of 10%. Underlying GDP per capita data is from the Penn World Table 7.0.
	Independent Variables	
$XCONST_0$	Constraints on the executive	From Polity IV data. Measures *de facto* independence of the executive. Scaled from 1 (no constraints) to 7 (fully constrained). Fixed at last year before slump.
ELF	Ethno-linguistic fractionalization	From Desmet et al. (2012), the original source is the Ethnologue data (15th edition). Cross-section.
ELF_0	Fractionalization of ethno-political groups	From Ethnic Power Relations data version 3.01 (Wimmer et al., 2009) supplemented with EPR-ETH 2.00 for small countries. Fixed at last year before slump.
POL	Ethno-linguistic polarization	From Desmet et al. (2012) using the Esteban and Ray (1994) measure with $\alpha = 1$ and $k = 4$. The original source is the Ethnologue data (15th edition). Cross-section.
POL_0	Ethno-political polarization	Computed using Ethnic Power Relations data version 3.01 (Wimmer et al., 2009) and Esteban and Ray (1994) measure with $\alpha = 1$ and $k = 4$. Fixed at last year before slump.
ELA_0	Asymmetries between ethno-political groups (relative to fractionalization at equal sizes).	Computed using Ethnic Power Relations data version 3.01 (Wimmer et al., 2009) supplemented with EPR-ETH 2.00 for small countries. Fixed at last year before slump.
$GROUPS_0$	Number of groups	———
$EGIPGRPS_0$	Number of included groups	———
$EXCLGRPS_0$	Number of excluded groups	———
$DOMPOP_0$	Dominant population (in %)	———
$MONPOP_0$	Monopoly population (in %)	———
	Control Variables	
GDP per capita	Log of initial real GDP per capita	From the Penn World Table 7.0. Fixed at last year before slump.
Regional dummies	Africa, Americas, Asia, Europe, and Oceania.	UN classification. Oceania is base.
Decade dummies	1950s, 1960s, 1970s, 1980s, 1990s, and 2000s.	Coded at beginning of slump. 2000s is base.

index developed by Esteban and Ray (1994):

$$POL_i \equiv k \sum_{j=1}^{J_i} \left(\frac{n_{ij}}{N_i}\right)^{1+\alpha} \left(1 - \frac{n_{ij}}{N_i}\right) \tag{10}$$

where $\alpha = 1$ (as they show in an auxiliary theorem) and $k = 4$ to scale the index between zero and one. Again, we use a version computed by Desmet et al. (2012) and one we compute for the EPR data (denoted POL_0).

While the polarization index captures the extent of bimodality of a distribution[10], it is not a measure of asymmetries (such as the existence of one large and many small groups). To capture these, we propose another simple measure of ethno-linguistic asymmetries:

$$ELA_i \equiv \frac{\sum_{j=1}^{J_i} \left(\frac{n_{ij}}{N_i}\right)^2 - \frac{1}{J_i}}{1 - \frac{1}{J_i}} = \frac{J_i}{J_i - 1} \left[\sum_{j=1}^{J_i} \left(\frac{n_{ij}}{N_i}\right)^2 - \frac{1}{J_i}\right], \ \forall \ J_i > 1 \tag{11}$$

and $ELA_i = 1$ if $J_i = 1$. The ELA index is simply the (normalized) difference between fractionalization with equal shares and observed fractionalization; it's a normalized Herfindahl index. We only compute this index for the EPR data (denoted ELA_0). Desmet et al. (2012) do not use this measure. Recall that for any particular number of groups, the ELF measure attains its maximum at an equal allocation. The global maximum is reached when, in the limit, each person constitutes an ethnic group. Contrary to the polarization or fractionalization measure, the ELA index is zero when the groups are of equal sizes and approaches unity as a single group becomes dominant. For the empirical analysis that follows, using the index of group asymmetries together with the number of groups allows us to analyze the effect of these two components of ethnic heterogeneity separately and investigate the more subtle aspects of the theoretical model.

We also obtain several additional variables from the EPR data. $GROUPS_0$ is the number of relevant (active) ethno-political groups. $EGIPGRPS_0$ is the number of included ethno-political groups at the last year before the slump; that is, groups with have some level access to executive power. $EXCLGRPS_0$ is the number of ethno-political groups without access to the political executive. Finally, $DOMPOP_0$ and $MONPOP_0$

[10] It attains its maximum at a symmetric bimodal distribution.

Table 3: Summary Statistics of Independent Variables

Variables	Obs	Mean	Std. Dev.	Min	Max
$XCONST_0$	58	3.48	2.49	1.00	7.00
ELF	58	45.39	33.71	0.07	95.98
ELF_0	57	36.00	25.71	0.00	80.39
POL	58	40.04	24.98	0.14	85.99
POL_0	57	19.35	16.42	0.00	56.95
ELA_0	57	48.75	33.40	0.10	100.00
$GROUPS_0$	57	4.19	6.43	0.00	47.00
$EGIPGRPS_0$	57	1.37	1.33	0.00	7.00
$EXCLGRPS_0$	57	2.33	6.17	0.00	46.00
$MONPOP_0$	57	0.21	0.36	0.00	0.97
$DOMPOP_0$	57	0.21	0.34	0.00	0.98
(Log) GDP per capita	58	8.53	1.21	5.87	10.63

are the population shares of the dominant or monopoly groups (the two highest levels of political power occurring only in mono-ethnic governments). All of these variables are fixed at the last year before the slump to rule out any feedback from the duration to group composition. Table 2 describes all variables and lists the underlying data sources. Table 3 presents the associated summary statistics.

Empirical approach. We employ standard event history techniques to study the duration of the decline phase. Survival analysis is particularly suitable for our purposes for two reasons. First, our empirical predictions are clearly about the time it take until the recovery starts and parametric duration models allow us to fully specify the underlying duration process, including the shape of the baseline hazard. Second, survival methods are designed to deal with censored observations which account for a non-trivial proportion of our sample. If the observed slump is completed, then the likelihood incorporates the information that the recovery has started at some point within the sample. Whereas if the slump is unfinished, only the fact that the country was still experiencing a decline enters the likelihood. Our approach is to examine partial correlations and test whether these are consistent with the proposed theory. While we cannot rule out all forms of endogeneity, we do take care to ensure temporal precedence by only linking pre-slump realizations to the subsequent duration. Note that in a duration set-up with few repeated spells, we cannot use country fixed-effects or a full set of time dummies (since time is

already parametrized) but we do include region dummies and decade dummies in most tables in the robustness section.

To estimate the partial correlations, we run log-normal accelerated failure time (AFT) regressions of the form:

$$\ln \tilde{t} \equiv \ln(t - t_0) = \beta_0 + \beta_1 XCONST_0 + \beta_2 H + \beta_3 (XCONST_0 \times H) + \mathbf{x}_0' \boldsymbol{\xi} + \epsilon_t \tag{12}$$

where $\tilde{t}$ is analysis time, t_0 is the last year before the slump, $XCONST_0$ is executive constraints, H is a measure of group (ethnic) heterogeneity, $\mathbf{x}_0$ is a vector of controls, and $\epsilon_t \sim \mathcal{N}(0, \sigma_\epsilon)$. Variables which could endogenously react to a prolonged duration of declines are kept fixed at t_0 to rule out any such feedback; if they have no time dimension, then we drop the subscript. All parameters are estimated using Maximum Likelihood and the standard errors are clustered on the country level to account for repeated spells.

Our main parameters of interest are β_1, β_2, and β_3. In several regressions, we impose $\beta_3 = 0$ to estimate first-order effects before examining the hypothesized interaction effect. The vector $\mathbf{x}_0$ may include additional heterogeneity measures, the (log of) initial GDP per capita before the slump, region effects and decade dummies.

Accelerated failure time models are so named due to their interpretation. A coefficient greater than zero implies that time passes more slowly, so that the exit of the decline phase is prolonged. A coefficient less than zero implies that time passes more quickly and hence that the recovery starts sooner. Alternatively, we may simply read the effects as elasticities (or semi-elasticities) of the expected duration with respect to the variables on the right hand side. Duration models have the main benefit of accounting for right censoring, otherwise their interpretation is identical to log-linear OLS when they are cast in the log-normal AFT form.

Results. Table 4 presents the first set of results corresponding mainly to the predictions derived from the baseline model. We compute two sets of estimates. One using the *Ethnologue* data which focuses on linguistic diversity and one based on the EPR data which incorporates only politically-relevant groups divided along the predominant social

cleavage (ethnic, linguistic, racial, etc.).

Columns (1) to (3) use the *Ethnologue* data. Column (1) establishes that stronger constraints on the political executive shorten the expected duration of the decline phase and that greater linguistic heterogeneity has an adverse effect on the expected duration. The effects are statistically significant at the 1%-level and economically meaningful. A one point improvement in executive constraints (before the slump) leads to an approximate 17.6% reduction in the duration until the trough. Conversely, a one percentage point change towards greater linguistic heterogeneity prolongs the decline phase by about 1.7%. Column (2) allows for a conditional effect and strongly suggests that the effect of political institutions depends on the level of linguistic diversity (and *vice versa*). Whenever we introduce an interaction term, we first center the two constituent variables on their average. This shifts the coefficients of the two base levels into a meaningful range, but leaves the magnitude and statistical significance of the interaction term unaffected. Holding the other variable constant, the coefficient on either base variable now measures the effect of a one unit change away from the average. As a result, the interaction effect can be ignored; it has to be taken into account only when both variables change. The interaction between executive constraints and linguistic fractionalization is significant at the 5%-level and comparatively large. The specification predicts that at perfect homogeneity the median decline lasts about 2 years, while at perfect heterogeneity it lasts about 12 years. These estimates cover nearly all of the observed differences between declines in Western Europe and Sub-Saharan Africa.

The results in columns (1) and (2) are consistent with our theoretical predictions; greater constraints on the executive shorten the expected duration unless the society is nearly homogeneous. The partial effect of executive constraints is not statistically different from zero for low ELF values. Column (3) adds the linguistic polarization measure to the specification in column (1). The literature on civil conflict stresses that polarization matters; e.g. Esteban and Ray (2011) show theoretically that conflict over public goods is driven by polarization and conflict over private goods by fractionalization. Contrary to this literature but in line with our model, we find no evidence in favor of

Table 4: Baseline – Executive Constraints, Heterogeneity and Interactions

	Dependent Variable: $\ln \tilde{t}$					
VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)
	Ethnologue			*Ethnic Power Relations*		
$XCONST_0$	-0.193***	-0.289***	-0.175***	-0.187***	-0.262***	-0.170**
	(0.060)	(0.084)	(0.062)	(0.067)	(0.085)	(0.067)
ELF	0.017***	0.019***	0.023***			
	(0.004)	(0.004)	(0.006)			
$XCONST_0 \times ELF$		-0.004**				
		(0.002)				
POL			-0.011			
			(0.007)			
ELF_0				0.020***	0.023***	0.025***
				(0.007)	(0.007)	(0.007)
$XCONST_0 \times ELF_0$					-0.004*	
					(0.002)	
POL_0						0.012
						(0.009)
	Control sets					
GDP per capita	Yes	Yes	Yes	Yes	Yes	Yes
	Summary stats					
Exits	48	48	48	47	47	47
Spells	58	58	58	57	57	57
Years of Decline	348	348	348	346	346	346
Log-$\mathcal{L}$	-74.704	-72.495	-73.645	-76.294	-74.952	-75.597
Pseudo-R^2	0.150	0.175	0.162	0.119	0.134	0.127

Note(s): The table shows the results from log-normal survival regressions of the duration of economic declines on our variables of interest. The standard errors are clustered on the country level to account for repeated spells. All specifications include a constant (not shown). Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

the hypothesis that polarization is an issue for (the lack of) cooperation during declines, while the coefficient on fractionalization is robust to this perturbation. In other words, existence of two equally powerful groups does not predict longer declines than, say, three equally-sized linguistic groups.

Measures of linguistic diversity tend to describe Sub-Saharan Africa as more diverse in comparison to other regions than alternative diversity measures. This begs the question if we are just estimating an "Africa effect". Columns (4) to (6) use the EPR data which addresses this issue by alternating the relevant cleavage by country (from racial over linguistic to religious). This changes the relative location of Sub-Saharan Africa, which

is only the second most diverse region on this measure, after South Asia, contrary to being the most linguistically diverse region based on the *Ethnologue* data. The EPR data also only codes politically relevant groups, so that the level of heterogeneity – no matter the measure – is generally lower. Note that we compute the heterogeneity measures for all politically relevant groups, not just the included groups. Strikingly, the results are virtually unchanged. Column (4) shows that the first order effects of executive constraints are the same, and the effect of fractionalization is well within one standard error of the previous estimate. Moreover, the sign and size of the interaction effect in column (5) is nearly identical to the one in column (2). Only the statistical significance of the interaction effect is a bit lower (cluster robust t-stat = -1.71). Column (5) shows that we also find no evidence in favor of *ethno-political* polarization affecting the duration of declines, just as with linguistic polarization. Contrasting these results to the *Ethnologue* data, it seems safe to conclude that we are not only explaining that declines in Sub-Saharan Africa last longer than elsewhere because the subcontinent is the most linguistically diverse. Our results also hold when we account for political relevance and vary the relevant divisions so that Sub-Saharan Africa is no longer the most diverse region in the world.

Overall, Table 4 provides significant evidence that there is a robust partial correlation of the duration of the decline phase with executive constraints on the one hand and with ethnic diversity on the other hand. In addition, the effect of weak constraints on the executive seems to be conditional on the degree of (ethno-political) fractionalization.

In Table 5 we "unpack" these statements further and examine what type of group configurations give rise to the cooperation problem we are analyzing. These results correspond mainly to the empirical content of the model extensions. We now only use the EPR data, as it provides the necessary detail on the number of groups, their power status and more. Column (1) presents a result that may appear puzzling at first sight. If we measure heterogeneity simply by the number of ethno-political groups, then we find no evidence of an effect on the duration of declines. However, this coefficient amalgamates two effects. The EPR data distinguishes between *included* groups, which have access

to executive power, and *excluded* groups, which lack power at the state level or are (at worst) discriminated against. In the model presented earlier, only the former are relevant players and thus we also expect that only they are empirically relevant. Column (2)

Table 5: Extensions – Number of Groups, Political Relevance, and Asymmetries

	Dependent Variable: $\ln \tilde{t}$					
VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)
$XCONST_0$	-0.225***	-0.241***	-0.215***	-0.179***	-0.210***	-0.196***
	(0.070)	(0.063)	(0.065)	(0.066)	(0.070)	(0.066)
$GROUPS_0$	-0.008			-0.031**		
	(0.018)			(0.014)		
$EGIPGRPS_0$		0.426***			0.290**	0.285**
		(0.095)			(0.124)	(0.113)
$EXCLGRPS_0$		-0.012			-0.021*	-0.013
		(0.013)			(0.012)	(0.012)
$DOMPOP_0$			-0.007*			
			(0.004)			
$MONPOP_0$			-0.011**			
			(0.005)			
ELF_0				0.022***	0.013	
				(0.007)	(0.009)	
ELA_0						-0.012**
						(0.005)
	Control sets					
GDP per capita	Yes	Yes	Yes	Yes	Yes	Yes
	Summary stats					
Exits	47	47	47	47	47	47
Spells	57	57	57	57	57	57
Years of Decline	346	346	346	346	346	334
Log-$\mathcal{L}$	-81.069	-75.062	-77.647	-75.457	-73.253	-71.791
Pseudo-R^2	0.064	0.133	0.103	0.129	0.154	0.171

Note(s): The table shows the results from log-normal survival regressions of the duration of economic declines on our variables of interest. The standard errors are clustered on the country level to account for repeated spells. All specifications include a constant (not shown). Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

confirms this expectation. The effect of included groups is statistically significant at the 1%-level and economically very large: an additional group increases the duration of the decline phase by about 53%. On the contrary, the effect of excluded groups is estimated to be near zero and has a comparatively tight 99% confidence interval centered near zero. In line with the theory, these results suggest that only ethnic groups with some degree of

access to political power matter for the duration of declines.

We have not yet isolated whether this adverse effect of heterogeneity is due to several equally powerful groups co-existing in the society or due to particular asymmetries in political power. Columns (3) to (6) represent different attempts towards empirically answering this question. In Comment 4, we translated the theoretical results regarding political concentration from Proposition 3 and the number of groups from Proposition 4 into the concepts of monopoly groups and dominant groups. Column (3) is the empirical counterpart. Here we relate the share of population represented by a group that either monopolizes or dominates the political executive to the duration of the decline phase. The results are unambiguous. Both variables are associated with substantially shorter declines. Through the lens of our framework, this finding is hardly surprising. The definition of dominant or monopoly groups means that they rule alone and are thus not bargaining over stabilization policies with other groups in the executive.

Columns (4) and (5) try to explicitly tackle the issue of the number of groups versus group asymmetries. In column (4), we include the number of ethno-politically relevant groups together with the index of ethno-political fractionalization. This leads to an interesting *ceteris paribus* condition. Increasing the degree of fractionalization by one percentage point while holding constant the number of groups necessarily implies that political concentration is decreasing; that is, the groups are becoming more alike. Recall that for any given number of groups, fractionalization is maximized at equal shares. The estimates thus suggest that less political concentration leads to longer declines. Column (5) again distinguishes between included and excluded groups to illustrate that only the former are relevant. The coefficient on the ELF_0 measure loses significance, suggesting that the number of included groups may drive the effect of ethnic heterogeneity and that group imbalances hardly matter. However, column (6) addresses this issue more directly by using our index of ethnic asymmetries and provides the same answer as column (4). Now the effect is easy to interpret, negative and significant at the 5% level. A one percentage point move towards greater asymmetries (political concentration) shortens the duration by about 1.3%. Note that the effect of executive constraints remains robust

throughout, fluctuating around a 20% reduction in the duration of declines for a one point improvement.

To summarize, Table 5 adds several valuable insights about the effect of ethnic diversity on the duration of declines. Fractionalization of linguistic or ethno-political groups masks two effects: 1) the expected duration is increasing in the number of politically relevant groups, and 2) the expected duration is decreasing in greater group asymmetries (political concentration). Both theory and evidence suggest that this is not an issue of polarization, but rather an issue of adding a smaller, potentially irrelevant, group to any multi-modal distribution of power as opposed to adding another powerful group (an additional mode).

Table 6 selects three key specifications using both data sources and then subjects them to two robustness checks. First, we return to the issue of whether we are estimating an "Africa effect" by including region dummies in each specification. Second, we control for temporal heterogeneity by including a dummy for the decade in which the slump began in every other specification, since the 1970s, 1980s and 1990s exhibit significantly higher volatility than the other decades.

The Africa dummy is significant in all specifications, capturing that declines take substantially longer on the African continent. Nevertheless, we still find comparable effects. Column (1) uses the *Ethnologue* data and shows that our two variables of interest are robustly correlated with within region differences in the duration of declines. Column (2) adds that this is still the case when we also control for temporal heterogeneity. Using the *EPR* data, columns (4) and (5) verify that the same holds for ethno-political fragmentation. The standard error of the interaction term becomes somewhat wider, leading to a loss of significance, but the estimated coefficient is extremely stable. The last two columns show that this also holds for the effect of the number of included groups. In general, there is significant evidence of regional heterogeneity (a χ^2-test always rejects the null of no heterogeneity at the 5%-level), but there is somewhat less evidence of temporal heterogeneity (on top of duration dependence).[11] Throughout Table 6 the coefficient of

[11] A χ^2-test rejects the null of no temporal heterogeneity at the 1%-level in column (1), at the 5%-level but not the 1%-level in column (4), and fails to reject the null at conventional levels in column (6).

Table 6: Robustness – Region and time effects

	Dependent Variable: $\ln \tilde{t}$					
VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)
$XCONST_0$	-0.256***	-0.251***	-0.211***	-0.171**	-0.181***	-0.145**
	(0.071)	(0.060)	(0.077)	(0.068)	(0.066)	(0.064)
ELF	0.020***	0.019***				
	(0.004)	(0.003)				
$XCONST_0 \times ELF$	-0.003***	-0.004***				
	(0.001)	(0.001)				
ELF_0			0.017**	0.015**		
			(0.007)	(0.006)		
$XCONST_0 \times ELF_0$			-0.003	-0.003*		
			(0.002)	(0.002)		
$EGIPGRPS_0$					0.298***	0.212*
					(0.097)	(0.119)
$EXCLGRPS_0$					0.024*	0.009
					(0.013)	(0.026)
Control sets						
GDP per capita	Yes	Yes	Yes	Yes	Yes	Yes
Regional dummies	Yes	Yes	Yes	Yes	Yes	Yes
Decade dummies	No	Yes	No	Yes	No	Yes
Summary stats						
Exits	48	48	47	47	47	47
Spells	58	58	57	57	57	57
Years of Decline	348	348	346	346	346	346
Log-$\mathcal{L}$	-63.635	-58.134	-67.966	-64.108	-68.705	-66.701
Pseudo-R^2	0.276	0.338	0.215	0.260	0.207	0.230

Note(s): The table shows the results from log-normal survival regressions of the duration of economic declines on our variables of interest. The standard errors are clustered on the country level to account for repeated spells. All specifications include a constant (not shown). Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

political institutions and the coefficients of the various measures of ethnic heterogeneity remain statistically significant at conventional levels and well within their usual range.

We report several further robustness checks in the appendix. Table A-2 uses a more lenient threshold for the identification of slumps (a significance level of 0.2). Our main results are qualitatively and quantitatively comparable in this larger set of 83 episodes. Table A-3 exchanges the fractionalization data with data on ethnic, linguistic and religious heterogeneity from Alesina et al. (2003), data on ethnic and cultural distance from Fearon (2003), and the original *Atlas Narodov Mira* data. For all but religious fractionalization, we find very similar interaction effects. In Table A-4, we include the Greenberg index,

which is a Gini index that accounts for linguistic distance, a peripheral heterogeneity index, which is a variant of the Greenberg index that accounts for the alienation between groups in the center and the periphery (both are from Desmet et al., 2009), genetic diversity from Ashraf and Galor (2013), and three measures of segregation from Alesina and Zhuravskaya (2011). Again, all interactions other than the one with religious segregation point in the right direction, while measures of *linguistic* heterogeneity tend to have the most robust effects.[12] This is in line with the extant empirical literature which tends to find that religious diversity often plays a different role than ethno-linguistic diversity (e.g. see Alesina et al., 2003; Alesina and Zhuravskaya, 2011). Next we examine if spatial and ethnic inequalities capture relevant notions of heterogeneity. We use a cross-section of spatial and ethnic inequalities in 1992 from Alesina et al. (2016) who estimate inequality based on differences in nighttime light intensities among arbitrary boxes, *Ethnologue* homelands at different levels of the linguistic tree, and homelands from the *Atlas Narodov Mira*. Table A-5 shows that while the coefficients on the constituent variables and the interaction term consistently point in the right direction, there is only weak evidence that ethnic inequality among linguistically distant groups (level 1) leads longer declines and interacts with political institutions. Table A-6 switches the Polity IV data with the political constraints data from Henisz (2000). Here too, the main results remain intact for the alternate measures of executive constraints. Finally, Table A-7 shows that the results do not depend on the specific functional form of the survival process.

Last but not least, we test a key assumption of our model. Throughout the paper we assume that *i)* weaker groups face an elevated risk of loosing political influence during a crisis and *ii)* that the risk of expropriation is particularly high when the recovery starts. For this test, we turn things around and examine if the number of included groups (denoted $EGIPGRPS_t$) deviates from its trend during the decline phase of the crisis

[12]We also broadened our concept of political heterogeneity to government fractionalization and legislative fractionalization. We then ran horse race regressions of these variables and their interactions with executive constraints, while keeping ethno-linguistic fractionalization, GDP per capita, region dummies, and time dummies in the specification. None of these variables or their interaction with executive constraints are significant at conventional levels, while the effect of ethno-linguistic fractionalization remains robust throughout.

Table 7: Number of politically-relevant groups

	Dependent Variable: $EGIPGRPS_t$					
VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)
Decline in t	-0.072**	-0.075**	-0.076**	-0.074**	-0.073**	-0.072**
	(0.031)	(0.031)	(0.031)	(0.031)	(0.032)	(0.032)
Trough in $t-1$	-0.119**					
	(0.058)					
Trough in $t-2$ to $t-1$		-0.093**				
		(0.042)				
Trough in $t-3$ to $t-1$			-0.067*			
			(0.035)			
Trough in $t-4$ to $t-1$				-0.043		
				(0.032)		
Trough in $t-5$ to $t-1$					-0.029	
					(0.029)	
Trough in $t-6$ to $t-1$						-0.024
						(0.028)
Control sets						
Country FEs	Yes	Yes	Yes	Yes	Yes	Yes
Times FEs	Yes	Yes	Yes	Yes	Yes	Yes
Country-specific Trends	Yes	Yes	Yes	Yes	Yes	Yes
Summary stats						
Within-R^2	0.426	0.426	0.426	0.425	0.425	0.425
N	51	51	51	51	51	51
$\bar{T}$	53.96	53.96	53.96	53.96	53.96	53.96
$N \times \bar{T}$	2752	2752	2752	2752	2752	2752

Note(s): The table shows the results from country-level panel regressions of the number of ethno-politically relevant groups on our indicators of the timing of economic slumps. All groups with a power status of 'junior partner' or higher are coded as politically relevant. The standard errors are clustered on the country level. All specifications include a constant (not shown). Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

and the immediate years after the trough. This implies that we are now dealing with a more traditional country-year panel over the period from 1950 to 2008. To purge most of the confounding heterogeneity, we always include country fixed-effects, time fixed-effects and country-specific linear time trends in the specifications.

Table 7 demonstrates that the data are consistent with our approach. First, the effect of a crisis on the number of included groups is always significant, suggesting that the assumed fear of falling out is real and intensifies during a downturn. Second, columns (1) to (3) show that this effect is also present and possibly even a bit larger in the

first years of the recovery.[13] Table A-8 confirm this finding using group-level regressions of the probability that a group is currently part of the executive on the same set of crisis indicators. In fact, the table shows that *individual* groups with junior or senior partner status in multi-ethnic governments face a statistically significant and non-trivial probability of falling out of the executive in the three years after the through (but do not tend to fall out during the decline phase). This adds another strong piece of evidence and suggests that our simplified focus on expropriation during recoveries has a basis in reality.

Taken together, these last empirical findings tell the following story. First, ethnic heterogeneity and constraints on the political executive are robust determinants of the length of the decline phase during economic slumps. Second, this result is not due to regional differences in ethnic heterogeneity but holds when we only use *within region* variation. Third, our main findings are robust to a variety of perturbations in the dependent and independent variables. Testing different notions of heterogeneity adds that our mechanism seems to be particularly relevant for countries that are politically divided along ethnic lines. Fourth, the dynamics of group participation in the executive are consistent with the assumptions and predictions of the model. Hence, we believe that our empirical approach operationalizes the model's key parameters and demonstrates that there is robust evidence in line with our theory.

4 Concluding remarks

This paper presents a political economy theory of declines. It highlights a commitment problem between winners and losers of the recovery process after a crisis, and then analyzes the empirical implications of this theory. We show that it is the combination of ethno-political heterogeneity with weak constraints on the political executive which brings about delayed cooperation. Together, these two factors help to explain why we observe such long declines in some countries and relatively short ones in others.

Both the theory and the empirical analysis suggest that ethnic heterogeneity is harmful

[13]Note that we cannot reject the hypothesis that that these two coefficients are in fact equal.

for getting groups to agree on a response to a crisis when political institutions are weak. More subtle predictions show that this is mostly an issue of having many powerful groups in the society and does not apply to the same degree when there is a politically dominant group. These findings have important policy implications. On the one hand, political institutions that constrain a country's leadership can contain the adversarial element of ethno-political heterogeneity. On the other hand, our research points out that there can be adverse effects of broad political inclusion when the institutional structure does not sufficiently limit executive power.

While not restricted to understanding declines in Sub-Saharan Africa, we would like to emphasize that we believe these insights are particularly important for understanding the political economy of declines on that subcontinent. Sub-Saharan Africa is home to the longest and deepest declines, politics shaped by ethnicity, and weak institutions governing executive power. While we still need to better understand why ethnic diversity tends to coincide with weak political institutions, we find that there is ample room for managing this heterogeneity better to avoid that welfare gains are lost again in the next crisis.

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For Online Publication: Appendix

Identifying slumps. In Bluhm et al. (2014) we outline a new approach to finding the duration of the decline phase of large economic slumps. Our restricted structural change approach is a variant of Bai (1997) and Papell and Prodan (2014).

We specify the following model for each (log) GDP per capita time series y_t:

$$y_t = \alpha + \beta t + \gamma_0 \mathbf{1}(t > tb_1) + \gamma_1 (t - tb_1)\mathbf{1}(t > tb_1) + \gamma_2 (t - tb_2)\mathbf{1}(t > tb_2) + \sum_{i=1}^{p} \delta_i y_{t-i} + \epsilon_t$$

where tb_1 and tb_2 are the endogenous break dates, $\mathbf{1}(\cdot)$ is an indicator function, and p is the lag order. The optimal $AR(p)$ model is determined by the Bayesian information criterion (BIC). We require that $tb_2 \geq tb_1 + 4$, so that the period between two successive breaks is at minimum 4 years.

We impose two restrictions. First, we require $\beta > 0$, so that growth must be positive in the years before a slump begins. Second, we also impose the condition that $\gamma_0 < 0$, so that a slump always starts with a drop in the intercept. Slope shifts are left unrestricted, so that the model can catch unfinished slumps (e.g., declines from tb_1 onwards, possibly lasting until the end of a country's time series). Next we compute the the sup-W test statistic of the null of no break versus two breaks ($\mathbb{H}_0 : \gamma_0 = \gamma_1 = \gamma_2 = 0$).

Note that the individual Wald tests over which the sup-W statistic is computed are not statistically independent. Hence, we bootstrap the empirical p-value of the sup-W statistic using a recursive bootstrap (Diebold and Chen, 1996). If the bootstrap test rejects at the desired significance level, α, we record the break pair $(\widehat{tb}_1, \widehat{tb}_2)$ and split the sample into a series running until the first break and a series starting just after the second break. The process starts again on each sub-sample until the bootstrap test fails to reject the null hypothesis of no breaks or the sample gets too small ($T < 20$). We set $\alpha = 0.1$ for the baseline results.

Next we date the though. For each of the identified episodes, we then define a censoring indicator $c = \mathbf{1}(\max_{j \in (\widehat{tb}_1, T]} y_j < y_{\widehat{tb}_1})$ that signifies if the slumps is finished or not. The end of a slump has occurred with certainty in the first year $a > \widehat{tb}_1$ where

$y_a \geq y_{\widehat{tb}_1}$; that is, when GDP per capita recovers to the level before the slump.

More formally, given the set of possible end years $A = \{a \mid a \in (\widehat{tb}_1, T] \text{ and } y_a \geq y_{\widehat{tb}_1}\}$, define $a_0 = \min A$ as the certain end of the slump. The estimated trough occurs at

$$t_{min} = \begin{cases} \text{argmin}_{j \in (\widehat{tb}_1, a_0]}\, y_j, & \text{if } c = 0 \\ \text{argmin}_{j \in (\widehat{tb}_1, T]}\, y_j, & \text{if } c = 1. \end{cases}$$

Last but not least, we define the duration of the decline phase as the duration of the beginning of the slump until the (provisional) trough, or $\tilde{t}_D = \hat{t}_{min} - \widehat{tb}_1$.

Proof of Proposition 1. The utility from cooperation in the first period when the other group cooperates is

$$v_j^1(C, c) = \frac{1}{1-\delta} \left\{(1-2c)\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] + c(g(0) + g(1))\right\} \tag{A-1}$$

and the utility from choosing to delay cooperation one period when the other group cooperates is

$$v_j^1(D, c) = g((1-\Delta)y_j) + \frac{\delta}{1-\delta}\left\{(1-2p^2)\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] + p^2(g(0) + g(1))\right\} \tag{A-2}$$

where $p^2 = c - x$; that is, half the probability of landing outside the safe zone in the second period.

The proof is by contradiction. We conjecture an equilibrium with immediate recovery, such that $v_j^1(C, c) \geq v_j^1(D, c)$. Using $p^2 = c - x$ and rearranging terms, we get

$$\begin{aligned} &g((1-\Delta)y_j) \leq \\ &\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] - \left[c + \frac{\delta}{1-\delta}x\right]\left\{2\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] - g(0) - g(1)\right\}. \end{aligned} \tag{A-3}$$

Note that concavity implies that $\{2\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] - g(0) - g(1)\} > 0$. Inequality (A-3) is contradicted whenever c, x or δ are large enough in relation to Δ, depending on the shape of the utility function $g(y_j)$ and its range, which completes the proof. ■

Proof of Proposition 2. First of all, it is useful to demonstrate that the difference in utility between recovery at any time period (t) and recovery at the subsequent period ($t+1$) decreases over time. For all $s > t$, we need to check whether

$$v_j^{t+1}(C,c) - v_j^t(C,c) > v_j^{s+1}(C,c) - v_j^s(C,c). \tag{A-4}$$

Note that $v_j^{t+1}(C,c) = v_j^t(D,c)$.

Substituting the utilities and rearranging the inequality, we get

$$\begin{aligned}
&g((1-\Delta)y_j) - (1-2p^t)\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] - p^t(g(0)+g(1)) + \\
&\frac{\delta}{1-\delta}\left\{2(p^t - p^{t+1})\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] + (p^{t+1} - p^t)(g(0)+g(1))\right\} > \\
&g((1-\Delta)y_j) - (1-2p^s)\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] - p^s(g(0)+g(1)) + \\
&\frac{\delta}{1-\delta}\left\{2(p^s - p^{s+1})\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] + (p^{s+1} - p^s)(g(0)+g(1))\right\}.
\end{aligned} \tag{A-5}$$

Recall that $p^t = c - (t-1)x$ implies $p^{t+1} - p^t = -x$, so the second and third terms cancel and the inequality reduces to

$$\begin{aligned}
&(1-2p^t)\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] + p^t(g(0)+g(1)) < \\
&(1-2p^s)\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] + p^s(g(0)+g(1)).
\end{aligned} \tag{A-6}$$

Substituting $p^t = c - (t-1)x$ again, it is straightforward to show that this inequality is always satisfied when $s > t$.

Having established this, setting the utility of choosing to cooperate in period t equal to the utility of recovering in period $t+1$ results in an equation that will deliver a potentially non-integer t, such that the smallest higher integer ($\lceil t \rceil$) is the equilibrium time to recovery:

$$\begin{aligned}
&\frac{1}{1-\delta}\left\{(1-2p^t)\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] + p^t(g(0)+g(1))\right\} = g((1-\Delta)y_j) + \\
&\frac{\delta}{1-\delta}\left\{(1-2p^{t+1})\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] + p^{t+1}(g(0)+g(1))\right\}.
\end{aligned} \tag{A-7}$$

Inserting the linear process on $p^t = c - (t-1)x$ yields

$$\begin{aligned}&\frac{1}{1-\delta}\left\{(1-2(c-(t-1)x))\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] + (c-(t-1)x)(g(0)+g(1))\right\} = \\ &g((1-\Delta)y_j) + \frac{\delta}{1-\delta}\left\{(1-2(c-tx))\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] + (c-tx)(g(0)+g(1))\right\}.\end{aligned} \tag{A-8}$$

Isolating the first term of the geometric series gives

$$\begin{aligned}&\left\{(1-2(c-(t-1)x))\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] + (c-(t-1)x)(g(0)+g(1))\right\} + \\ &\frac{\delta}{1-\delta}\left\{(1-2(c-(t-1)x))\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] + (c-(t-1)x)(g(0)+g(1))\right\} = \\ &g((1-\Delta)y_j) + \frac{\delta}{1-\delta}\left\{(1-2(c-tx))\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] + (c-tx)(g(0)+g(1))\right\}\end{aligned} \tag{A-9}$$

and after canceling the common terms, we have

$$\begin{aligned}&(1-2c)\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] + c(g(0)+g(1)) + \\ &tx\{2\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] - (g(0)+g(1))\} \\ &= g((1-\Delta)y_j) + \frac{1}{1-\delta}x\left\{2\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] - (g(0)+g(1))\right\}.\end{aligned} \tag{A-10}$$

Solving for t^* and simplifying gives

$$t^* = \frac{g((1-\Delta)y_j) - \mathbb{E}[g(w_j)|w_j \in \mathcal{A}]}{x\{2\mathbb{E}[g(w_j)|w_j \in \mathcal{A}] - (g(0)+g(1))\}} + \frac{c}{x} + \frac{1}{1-\delta}. \tag{A-11}$$

The proposition follows directly from comparative statics w.r.t. to c

$$\frac{\partial t^*}{\partial c} = \frac{1}{x} > 0 \tag{A-12}$$

that is, stronger executive constraints (smaller c) shorten the time to cooperation. This completes the proof for the interior case. Note that it can also be the case that recovery happens at the point when all uncertainty is resolved, i.e. the point where the probability of being outside the safe zone is zero and no longer changes. If this is the case it is straightforward to see that the time to recovery is shorter with stronger constraints on the executive. This follows directly from the fact, that the time it takes until all uncertainty

is resolved is shorter with smaller c. ■

Proof of Proposition 3. Recall that for the asymmetric case only one group risks falling outside the political safe zone. Hence, for there to exist an equilibrium with recovery in the first period, the following condition needs to be true

$$\frac{1}{1-\delta}\left\{(1-p_1(y_j))y_j + p_1(y_j)z\right\} \geq (1-\Delta)y_j + \frac{\delta}{1-\delta}y_j \tag{A-13}$$

which simplifies to

$$\Delta y_j + \frac{1}{1-\delta}\left\{p_1(y_j)(z-y_j)\right\} \geq 0. \tag{A-14}$$

An decrease in concentration (asymmetry) makes delay more likely if the left hand side of the inequality is a decreasing function of y_j. This is true when the derivative of the left hand side is negative:

$$\Delta + \frac{1}{1-\delta}\left\{\frac{dp_1(y_j)}{dy_j}(z-y_j) - p_1(y_j)\right\} < 0 \tag{A-15}$$

which completes the proof. ■

Proof of Proposition 4. Remember that there always exists an equilibrium with recovery in period two in the sub-game that starts in period two after delay in period one. If all of the other groups decide to cooperate in period 1, it is optimal for the remaining group to cooperate if the following condition holds

$$\begin{aligned}&\frac{1}{1-\delta}\left\{(1-Jp_1(y_j))\frac{1}{J} + (J-1)p_1(y_j)\left(\frac{1}{J} + \frac{1}{(J-1)J}\right) + p_1(y_j)z\right\} \geq \\ &(1-\Delta)\frac{1}{J} + \frac{\delta}{1-\delta}\frac{1}{J}.\end{aligned} \tag{A-16}$$

The second term inside the curly braces simplifies to $p_t(y_j)$, so that inequality (A-16) becomes

$$\frac{1}{1-\delta}\left\{(1-Jp_1(y_j))\frac{1}{J}+p_1(y_j)+p_1(y_j)z\right\} \geq (1-\Delta)\frac{1}{J}+\frac{\delta}{1-\delta}\frac{1}{J} \tag{A-17}$$

or

$$\frac{1}{1-\delta}\left\{\frac{1}{J}+p_1(y_j)z\right\} \geq (1-\Delta)\frac{1}{J}+\frac{\delta}{1-\delta}\frac{1}{J} \tag{A-18}$$

and, after some algebraic manipulation, this simplifies to

$$\frac{\Delta}{J}+\frac{1}{1-\delta}p_1(y_j)z \geq 0. \tag{A-19}$$

Now remember that $p_1(y_j)$ is increasing in J, as symmetry implies $y_j = 1/J$ and $z < 0$. As a result, the inequality becomes harder to satisfy if the number of groups increases, proving part *i)* of Proposition 4.

For part *ii)*, consider first a setting with $c = 0$. Then $p_1(y_j) = 0$ for all j and delay will not occur. Now introduce $c > 0$; in this case $p_1(y_j) \geq p_1(y_i)$ for all $j > i$ by assumption, and from part *i)* of this proposition we know that an increase in the number of groups makes delay more likely. Hence, the introduction of a threshold c greater than zero is more likely to imply delay when the number of groups is larger. ■

Table A-1: Estimated breaks with troughs: 58 episodes

Code	T_0	$\widehat{tb}_1$	$\widehat{t}_{min}$	$\widehat{tb}_2$	T	Sup-W	Critical W	p-value	Drop (%)	Duration
ALB	1970	1990	1991	2002	2008	18.5	13.6	0.007	-15.32	1
ARE	1986	1990	1999	2002	2008	29.1	14.5	0.003	-10.90	9
AUS	1950	1954	1957	1966	2008	8.3	8.7	0.064	-0.72	3
AUS	1967	1989	1991	1998	2008	10.1	10.7	0.059	-2.29	2
BDI	1960	1971	1972	1988	2008	9.9	11.3	0.089	-3.23	1
BEL	1950	1957	1958	1973	2008	12.8	12.1	0.029	-2.24	1
BGR	1970	1988	1997	1997	2008	16.3	12.8	0.010	-23.79	9
BHR	1970	1980	1987	1986	2008	14.4	11.0	0.010	-44.12	7
BRA	1950	1980	1983	2003	2008	12.5	12.3	0.043	-14.60	3
CAF	1960	1978	2005	2005	2008	8.3	8.7	0.060	-46.38	27
CHE	1950	1974	1975	1978	2008	10.7	10.6	0.047	-7.87	1
CHL	1951	1953	1954	1972	1973	12.0	8.5	0.017	-9.06	1
CHL	1951	1974	1975	1979	1980	13.3	10.8	0.021	-16.50	1
CHL	1951	1981	1983	1995	2008	12.6	11.4	0.025	-21.22	2
CHN	1952	1960	1962	1977	2008	13.9	12.9	0.029	-23.71	2
CMR	1960	1986	1995	1990	2008	12.0	12.3	0.055	-40.46	9
COG	1960	1974	1977	1982	2008	11.9	12.5	0.069	-21.35	3
CRI	1950	1955	1956	1963	1979	11.4	11.3	0.048	-4.39	1
CRI	1950	1980	1982	2002	2008	17.2	10.6	0.002	-17.47	2
CUB	1970	1988	1993	1995	2008	11.4	12.5	0.072	-34.70	5
CYP	1950	1973	1975	1977	2008	15.5	9.7	0.001	-31.40	2
CYP	1978	1990	1991	1995	2008	11.6	14.6	0.098	-10.19	1
DNK	1950	1954	1955	1965	2008	12.9	11.7	0.022	-1.56	1
DZA	1960	1984	1994	1996	2008	10.9	8.2	0.013	-14.09	10
ETH	1950	1972	1992	1993	2008	11.5	10.2	0.020	-30.68	20
FIN	1950	1989	1993	2006	2008	10.6	10.8	0.057	-16.34	4
GAB	1960	1976	1987	1997	2008	10.6	11.2	0.062	-50.56	11
GMB	1960	1982	1998	2002	2008	16.4	11.2	0.006	-25.33	16
GRC	1951	1973	1974	1994	2008	17.9	11.6	0.003	-6.92	1
GTM	1950	1980	1988	1984	2008	15.1	12.3	0.015	-19.14	8
HUN	1970	1990	1992	2004	2008	15.6	13.5	0.018	-10.56	2
IDN	1960	1997	1999	2001	2008	13.5	10.6	0.013	-17.49	2
IRN	1955	1976	1981	1980	2008	15.9	11.6	0.004	-56.78	5
IRQ	1970	1990	2003	1994	2008	9.1	8.9	0.046	-66.43	13
JPN	1950	1973	1974	1990	2008	13.5	13.4	0.050	-2.85	1
MEX	1950	1981	1988	1995	2008	11.9	11.0	0.038	-17.03	7
MNG	1970	1990	1993	2003	2008	46.5	11.7	0.000	-41.81	3
MOZ	1960	1981	1986	1995	2008	12.6	12.0	0.037	-24.99	5
MYS	1955	1984	1986	1993	2008	9.1	10.5	0.093	-7.47	2
NPL	1960	1979	1980	2000	2008	10.6	8.9	0.025	-5.33	1
NZL	1950	1974	1978	1992	2008	9.9	10.5	0.070	-9.03	4
OMN	1970	1979	1980	1985	2008	12.4	9.0	0.007	-21.61	1
PER	1950	1958	1959	1966	1976	11.9	9.3	0.022	-6.91	1
PER	1950	1977	1992	1992	2008	11.0	10.3	0.037	-29.30	15
PHL	1950	1983	1985	2003	2008	12.8	10.2	0.007	-16.78	2
POL	1970	1979	1982	1993	2008	13.8	12.1	0.027	-22.55	3
PRY	1980	1989	2002	2002	2008	8.8	8.8	0.049	-14.24	13

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Table A-1 – *Continued from previous page*

Code	T_0	$\widehat{tb}_1$	$\widehat{t}_{min}$	$\widehat{tb}_2$	T	Sup-W	Critical W	p-value	Drop (%)	Duration
RWA	1960	1993	1994	1997	2008	18.0	7.9	0.001	-45.38	1
SAU	1986	1992	1999	2002	2008	14.6	13.3	0.039	-18.75	7
SLE	1961	1995	1999	2006	2008	14.2	11.1	0.011	-41.65	4
SLV	1950	1978	1983	1987	2008	18.2	10.2	0.002	-25.82	5
TGO	1960	1979	2008	1989	2008	9.6	10.1	0.065	-53.60	29
THA	1950	1996	1998	2003	2008	10.7	7.8	0.003	-14.17	2
TTO	1950	1961	1963	1969	1981	16.8	14.9	0.020	-0.78	2
TTO	1950	1982	1993	2006	2008	12.4	12.6	0.054	-28.96	11
UGA	1950	1977	1986	1987	2008	11.6	10.5	0.029	-30.27	9
USA	1950	1957	1958	1966	2008	8.7	9.3	0.075	-2.51	1
ZMB	1955	1968	2001	2000	2008	15.0	10.9	0.007	-68.99	33

Table A-2: Robustness – Sample of Slumps

	Dependent Variable: $\ln \tilde{t}$					
VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)
	Ethnologue			*Ethnic Power Relations*		
$XCONST_0$	-0.195***	-0.245***	-0.187***	-0.180***	-0.220***	-0.173***
	(0.050)	(0.054)	(0.051)	(0.054)	(0.051)	(0.053)
ELF	0.013***	0.014***	0.017***			
	(0.004)	(0.003)	(0.004)			
$XCONST_0 \times ELF$		-0.003**				
		(0.001)				
POL			-0.007			
			(0.006)			
ELF_0				0.013**	0.014***	0.015***
				(0.005)	(0.005)	(0.005)
$XCONST_0 \times ELF_0$					-0.003**	
					(0.001)	
POL_0						0.008
						(0.008)
	Control sets					
GDP per capita	Yes	Yes	Yes	Yes	Yes	Yes
	Summary stats					
Exits	71	71	71	69	69	69
Spells	83	83	83	81	81	81
Years of Decline	468	468	468	464	464	464
Log-$\mathcal{L}$	-114.133	-111.929	-113.498	-114.993	-113.526	-114.556
Pseudo-R^2	0.093	0.110	0.098	0.068	0.080	0.072

Note(s): The table shows the results from log-normal survival regressions of the duration of economic declines on our variables of interest. The larger sample of slumps was obtained by running the structural break algorithms and dating the trough using a more lenient significance threshold of $\alpha = .20$. The standard errors are clustered on the country level to account for repeated spells. All specifications include a constant (not shown). Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table A-3: Robustness – Measures of Fractionalization

	Dependent Variable: $\ln \tilde{t}$					
VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)
	Alesina et al.			*Fearon*		*Atlas*
$XCONST_0$	-0.230***	-0.293***	-0.184**	-0.239***	-0.243***	-0.284***
	(0.069)	(0.082)	(0.073)	(0.073)	(0.053)	(0.073)
Ethnic (H^A)	0.020***					
	(0.006)					
$XCONST_0 \times H^A$	-0.004**					
	(0.002)					
Linguistic (H^B)		0.021***				
		(0.006)				
$XCONST_0 \times H^B$		-0.004***				
		(0.002)				
Religious (H^C)			0.005			
			(0.008)			
$XCONST_0 \times H^C$			-0.004*			
			(0.002)			
Ethnic (H^D)				0.019***		
				(0.006)		
$XCONST_0 \times H^D$				-0.005***		
				(0.002)		
Cultural (H^E)					0.028***	
					(0.005)	
$XCONST_0 \times H^E$					-0.008***	
					(0.002)	
Ethnic (H^F)						0.020***
						(0.005)
$XCONST_0 \times H^F$						-0.005***
						(0.002)
	Control sets					
GDP per capita	Yes	Yes	Yes	Yes	Yes	Yes
Regional dummies	Yes	Yes	Yes	Yes	Yes	Yes
Decade dummies	Yes	Yes	Yes	Yes	Yes	Yes
	Summary stats					
Exits	48	45	48	48	48	45
Spells	58	55	58	58	58	55
Years of Decline	348	337	348	348	348	333
Log-$\mathcal{L}$	-63.681	-55.225	-67.932	-63.073	-58.602	-58.670
Pseudo-R^2	0.275	0.341	0.227	0.282	0.333	0.298

Note(s): The table shows the results from log-normal survival regressions of the duration of economic declines on our variables of interest. The measures of heterogeneity H^A to H^C are obtained from Alesina et al. (2003), H^D and H^E are from Fearon (2003), and H^F is from the Atlas Narodov Mira as published by Alesina et al. (2003). The standard errors are clustered on the country level to account for repeated spells. All specifications include a constant (not shown). Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table A-4: Robustness – Alternate Measures of Heterogeneity

	Dependent Variable: $\ln \tilde{t}$					
VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)
	Desmet et al.		*Ashraf & Galor*	*Alesina & Zhuravskaya*		
$XCONST_0$	-0.178***	-0.164***	-0.144**	-0.235**	-0.337***	-0.066
	(0.049)	(0.049)	(0.065)	(0.111)	(0.102)	(0.167)
Greenberg (H^A)	0.027***					
	(0.008)					
$XCONST_0 \times H^A$	-0.008***					
	(0.003)					
Peripheral Het. (H^B)		0.033**				
		(0.013)				
$XCONST_0 \times H^B$		-0.011***				
		(0.004)				
Genetic Div. (H^C)			0.065			
			(0.066)			
$XCONST_0 \times H^C$			-0.002			
			(0.009)			
Ethnic Seg. (H^D)				0.033		
				(0.021)		
$XCONST_0 \times H^D$				-0.011		
				(0.009)		
Linguistic Seg. (H^E)					0.032**	
					(0.015)	
$XCONST_0 \times H^E$					-0.014**	
					(0.007)	
Religious Seg. (H^F)						0.049**
						(0.024)
$XCONST_0 \times H^F$						0.019
						(0.031)
	Control sets					
GDP per capita	Yes	Yes	Yes	Yes	Yes	Yes
Regional dummies	Yes	Yes	Yes	Yes	Yes	Yes
Decade dummies	Yes	Yes	Yes	Yes	Yes	Yes
	Summary stats					
Exits	48	48	48	31	32	25
Spells	58	58	58	38	38	32
Years of Decline	348	348	348	249	248	235
Log-$\mathcal{L}$	-62.468	-63.393	-68.777	-39.815	-38.150	-34.907
Pseudo-R^2	0.289	0.278	0.217	0.342	0.366	0.307

Note(s): The table shows the results from log-normal survival regressions of the duration of economic declines on our variables of interest. The measures of linguistic heterogeneity H^A and H^B are obtained from Desmet et al. (2009), genetic diversity H^C is from Ashraf and Galor (2013), and the segregation measures H^D to H^F are from Alesina and Zhuravskaya (2011). The standard errors are clustered on the country level to account for repeated spells. All specifications include a constant (not shown). Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table A-5: Robustness – Measures of Spatial and Ethnic Inequality

	Dependent Variable: $\ln \tilde{t}$					
VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)
$XCONST_0$	-0.148**	-0.158**	-0.157**	-0.186***	-0.181**	-0.165**
	(0.072)	(0.065)	(0.063)	(0.070)	(0.073)	(0.077)
Spatial (H^A)	0.005					
	(0.008)					
$XCONST_0 \times H^A$	-0.001					
	(0.002)					
Ethnic Level 1 (H^B)		0.012**				
		(0.005)				
$XCONST_0 \times H^B$		-0.004*				
		(0.002)				
Ethnic Level 5 (H^C)			0.007			
			(0.005)			
$XCONST_0 \times H^C$			-0.002			
			(0.001)			
Ethnic Level 10 (H^D)				0.008*		
				(0.005)		
$XCONST_0 \times H^D$				-0.003*		
				(0.001)		
Ethnic Level 15 (H^E)					0.007	
					(0.005)	
$XCONST_0 \times H^E$					-0.002	
					(0.001)	
Ethnic GREG (H^F)						0.014*
						(0.007)
$XCONST_0 \times H^F$						-0.002
						(0.002)
Control sets						
GDP per capita	Yes	Yes	Yes	Yes	Yes	Yes
Regional dummies	Yes	Yes	Yes	Yes	Yes	Yes
Decade dummies	Yes	Yes	Yes	Yes	Yes	Yes
Summary stats						
Exits	48	48	48	48	48	48
Spells	58	58	58	58	58	58
Years of Decline	348	348	348	348	348	348
Log-$\mathcal{L}$	-68.875	-65.912	-67.546	-66.854	-67.546	-66.702
Pseudo-R^2	0.216	0.250	0.231	0.239	0.231	0.241

Note(s): The table shows the results from log-normal survival regressions of the duration of economic declines on our variables of interest. The measures of spatial and ethnic inequality H^A to H^F are obtained from Alesina et al. (2016). The standard errors are clustered on the country level to account for repeated spells. All specifications include a constant (not shown). Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table A-6: Robustness – Measures of Political Constraints

	Dependent Variable: $\ln \tilde{t}$					
VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)
	Henisz Political Constraints Data					
ELF	0.017***	0.020***	0.021***	0.021***	0.022***	0.022***
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
POLCON III	-1.317**	-2.130***				
	(0.568)	(0.631)				
POLCON III × *ELF*		-0.041***				
		(0.016)				
POLCON V			-0.901**	-1.092**		
			(0.399)	(0.477)		
POLCON V × *ELF*				-0.009		
				(0.010)		
POLCON VJ					-1.076	-2.289**
					(0.730)	(1.019)
POLCON VJ × *ELF*						-0.027*
						(0.016)
	Control sets					
GDP per capita	Yes	Yes	Yes	Yes	Yes	Yes
Regional dummies	Yes	Yes	Yes	Yes	Yes	Yes
Decade dummies	Yes	Yes	Yes	Yes	Yes	Yes
	Summary stats					
Exits	47	47	39	39	34	34
Spells	57	57	49	49	44	44
Years of Decline	347	347	335	335	325	325
Log-$\mathcal{L}$	-62.983	-60.722	-50.894	-50.602	-45.363	-44.934
Pseudo-R^2	0.269	0.295	0.304	0.308	0.305	0.311

Note(s): The table shows the results from log-normal survival regressions of the duration of economic declines on our variables of interest. Executive constraints are measured using the data from Henisz (2000). *POLCONIII* is derived from a structural veto-player model. *POLCONV* adds two additional veto points for the judiciary and sub-federal entities. *POLCONVJ* includes measures of alignment and fractionalization of the High Court. Only *POLCONIII* is still remotely related to the parameter c in our model. However, these measures always include legislative fractionalization, while we are concerned with ethnic fractionalizaton of the executive. The standard errors are clustered on the country level to account for repeated spells. All specifications include a constant (not shown). Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table A-7: Robustness – Functional Form

	Dependent Variable: $\ln \tilde{t}$					
VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)
	Coefficients ($\mathbb{H}_0 = 0$)		*Hazard Ratios* ($\mathbb{H}_0 = 1$)			
	Log-logistic		*Weibull*		*Cox*	
$XCONST_0$	-0.270***	-0.253***	1.455***	1.547***	1.323***	1.352***
	(0.084)	(0.075)	(0.126)	(0.147)	(0.099)	(0.097)
ELF	0.020***	0.020***	0.968***	0.963***	0.976***	0.974***
	(0.004)	(0.004)	(0.008)	(0.009)	(0.006)	(0.006)
$XCONST_0 \times ELF$	-0.004**	-0.004***	1.004*	1.008***	1.004**	1.005***
	(0.001)	(0.001)	(0.002)	(0.003)	(0.002)	(0.002)
	Control sets					
GDP per capita	Yes	Yes	Yes	Yes	Yes	Yes
Regional dummies	Yes	Yes	Yes	Yes	Yes	Yes
Decade dummies	No	Yes	No	Yes	No	Yes
	Summary stats					
Exits	48	48	48	48	48	48
Spells	58	58	58	58	58	58
Years of Decline	348	348	348	348	348	348
Log-$\mathcal{L}$	-64.578	-59.569	-66.091	-59.615	-148.403	-145.124
Pseudo-R^2	0.274	0.330	0.306	0.374	0.103	0.123

Note(s): The table shows the results from survival regressions of the duration of economic declines on our variables of interest. Columns (1) and (2) are AFT models with a log-logistic density, columns (3) and (4) are hazard models with a Weibull hazard, and columns (5) and (6) are non-parametric proportional hazard models. The standard errors are clustered on the country level to account for repeated spells. All specifications include a constant (not shown). Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table A-8: Robustness – Ethno-political relevance of groups in plurality-rule governments

	Dependent Variable: $Pr(EGIP)_{git}$					
VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)
Decline in t	0.004	0.002	0.000	-0.001	-0.002	-0.002
	(0.022)	(0.023)	(0.022)	(0.022)	(0.022)	(0.022)
Trough in $t-1$	-0.055**					
	(0.024)					
Trough in $t-2$ to $t-1$		-0.044*				
		(0.022)				
Trough in $t-3$ to $t-1$			-0.037*			
			(0.021)			
Trough in $t-4$ to $t-1$				-0.032		
				(0.021)		
Trough in $t-5$ to $t-1$					-0.030	
					(0.020)	
Trough in $t-6$ to $t-1$						-0.026
						(0.021)
	Control sets					
Country FEs	Yes	Yes	Yes	Yes	Yes	Yes
Times FEs	Yes	Yes	Yes	Yes	Yes	Yes
Country-Year Trends	Yes	Yes	Yes	Yes	Yes	Yes
	Summary stats					
Adjusted-R^2	0.678	0.678	0.678	0.678	0.678	0.678
G	150	150	150	150	150	150
$\bar{T}$	36.67	36.67	36.67	36.67	36.67	36.67
$G \times \bar{T}$	5500	5500	5500	5500	5500	5500

Note(s): The table shows the results from linear group-level panel regressions of the probability of being a politically relevant group with government participation on our indicators of the timing of economic slumps. The group-level data is from the EPR-ETH 2.00 (which includes small countries missing in EPR 3.01). The dependent variable is called `status_egip` in the original data. All groups with a power status of 'junior partner' or higher are coded as unity, all others as zero. We drop countries ruled by dominant and monopoly groups as these are single-ethnicity governments which are not the object of interest in our theory. The standard errors are clustered on the country level. Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.