Riots and the Window of Opportunity for Coup Plotters

Evidence on the Link between Urban Protests and Coups d'État

Lena Gerling*

Münster University, Center for Interdisciplinary Economics,
Scharnhorststrasse 100, 48151 Münster, Germany

November 2016

This paper investigates the impact of urban protests on coup attempts in a sample of 39 Sub-Saharan African countries for the period 1990 to 2007. Widespread public discontent, especially when occurring in urban centers, can act as a trigger of coups d'état in autocratic regimes by opening a window of opportunity for leadership removals by the ruling elite. The main difficulty in testing this relationship is that public revolts are rarely exogenous to coup risk. To address this problem, variation in rainfall is used to create an instrument for urban protests. The results show that rainfall-related popular uprisings in urban areas increase the likelihood of a coup attempt and thus help to solve the collective-action problems associated with coup plots.

JEL classification: C26, D74, P16

Keywords: Coup d'état, public protest, regime change, autocracy.

* *Contact information:* `lena.gerling@uni-muenster.de`; Tel.:+49 251 24304

1 Introduction

This paper departs from the notion that autocracies are political systems in which the government is appointed to or removed from power not by regular competitive elections but by force (or threats of force). Two common forms of force in autocracies are coups d'état that arise from a (formerly) loyal elite surrounding the incumbent government and mass movements of public protest. They frequently occur in autocratic regimes and challenge authoritarian rule. While the root cause of coup risk has been indentified in the level of economic development and regime legitimacy (Belkin and Schofer 2003; Londregan and Poole 1990), widespread public discontent seems to be a trigger of actual coup attempts as happened most recently in the "Arab Spring" countries Egypt and Tunesia, and in Thailand in 2014. Thus, public unrest, especially when occurring in urban centers, might open a window of opportunity for violent leadership removal by the elite. This paper empirically investigates the impact of urban protests on coup attempts in a sample of 39 Sub-Saharan African (SSA) countries for the period 1990 to 2007 by isolating exogenous variation in protests through rainfall variation.

Coordination among elite members is costly in autocracies as individuals planning to overthrow a current leader must fear repression and persecution if a conspiracy is discovered (Powell 2012). Caspar and Tyson (2014) suggest that "popular protests provide an overt statement of government illegitimacy" (p.548) because they signal discontent with the current leadership and weak repressive capacity. The publicity of this signal improves the individual's belief about the actions of other elite members, thereby enhancing coordination and thus the ability to stage a successful coup. In addition, public protests also increase the disposition of elites to intervene, either because coup plotters see an opportunity to revise policies in their respective favor (Apolte 2015), or because plotters feel the need to launch a coup in the presence of a "revolutionary threat" imposed by protesters that risk the elite's social status in the future (Acemoglu and Robinson 2006; Gilli and Li 2015).

The link between protests and coups has also been established in empirical studies (Belkin and Schofer 2003). Applying extreme bound analysis, Gassebner, Gutmann, and Voigt (2016) find that government crises, political instability and riots are among the most robust predictors of coup attempts. These results mirror earlier studies showing that public uprisings are positively associated with coups (Belkin and Schofer 2003; Thyne 2010; Powell 2012). Yet, the main difficulty in testing whether public protests solve the collective-action problems of the elite is that public revolts are rarely exogenous to coup risk. There might be reversed causality because the expectation of a coup plot itself can lead to riots. Moreover, unobserved variables could drive both, public protests and coup attempts. We contribute to the existing literature by addressing potential endogeneity problems with the use of exogenous variation in rainfall to create an instrumental variable (IV) for public protests.[1] To test the hypothesis that popular unrest increases the likelihood of a coup attempt, we focus on anti-government protests (demonstrations, strikes and riots) that are located in urban areas in SSA countries. There are good theoretical reasons to focus on urban protests as an important source of changing elite perceptions because urban dwellers are geographically concentrated and thus more likely to engage in collective action and to attract the attention of political elites than citizens in sparsely populated areas (Aidt and Leon 2015; Hendrix and Haggard 2015). Johnson and Thyne (2016) show that the strength of signals sent by public unrest to coup plotters is most credible when protests are located in urban areas. We use droughts (i.e. extreme negative rainfall deviations) as instrument for urban unrest, since there is robust evidence that adverse weather shocks trigger social unrest in countries that are dependent on rain-fed agriculture (e.g. Miguel, Satyanath, and Sergenti 2004; Hendrix and Salehyan 2012; Aidt and Leon 2015). Moreover, recent evidence suggests that weather shocks, through their impact on local food prices, increases the level of conflict in urban areas (Smith 2014; Raleigh, Choi, and Kniveton 2015). We also consider potential spatial heterogeneity in the link between adverse weather shocks and urban unrest and interact our drought instrument with region dummies for Western, Eastern, Middle and Southern Africa.

The IV estimates show a strong first stage relationship between variation in rainfall and the incidence of urban protests. The results suggest that a drought leads to an increase in the likelihood of a coup attempt of 5.49 percentage points through the increased duration of urban protests. This corresponds to a doubling in the average coup risk during drought years. Overall, the evidence provided in this paper yields strong and robust support to the idea that drought-related urban uprisings trigger coup attempts by opening a window of opportunity for ruling elites to overthrow an existing leadership.

The empirical strategy is most closely related to Aidt and Leon (2015). In their study, the authors analyze the effect of riots on political transitions in SSA countries, instrumenting riots by droughts. While our empirical procedure is similar to their approach, the dependent variable of interest in our paper relates to a different type of political

[1] We use the terms public protests, unrest and uprisings interchangeably here and distinguish these lower-scale forms of conflict from large-scale civil wars.

change, namely coerced leadership removal within the ruling class. Moreover, we focus on urban protests and allow for regional heterogeneity in the relationship between droughts and social unrest. Yet, both studies contribute to a better understanding of the factors driving political change. Our work also relates to the empirical literature on the determinants of coup attempts and aims at contributing to this field by disentangling more closely the causal links between public uprisings and elite decisions to defect (Casper and Tyson 2014; Johnson and Thyne 2016). Finally, our paper expands work on the relationship between weather shocks and low-scale social conflicts by allowing for regional heterogeneity of the impact of droughts on protests (Miguel, Satyanath, and Sergenti 2004; Burke, Hsiang, and Miguel 2014).

The remainder of the paper of the paper is organized as follows: Section two gives a short overview of the related theoretical and empirical literature. Section three explains the data sources and measurement of variables, and section four introduces the empirical strategy applied. Section five discusses our main results, while section six presents some robustness tests for our baseline findings. Finally, section seven points out potential caveats of the chosen approach and concludes.

2 Literature Review

Theoretical Links between Protests and Coups

Generally, regime elites face severe collection-action problems when organizing a coup plot. Coordination among elite members is costly in autocracies as individuals planning to overthrow a current leader must fear repression and persecution if a conspiracy is discovered (Powell 2012). Since information about the degree of loyalty among the supporting elite members is not public, the evolution of an effective coup conspiracy depends on the elite members' beliefs about other individuals' actions and loyalty. In addition to this strategic uncertainty associated with a coup plot, Bueno de Mesquita et al. (2005) show that elite members might be caught in a "loyalty trap" that binds them to the incumbent. The reason is that a potential challenger is not able to commit to a credible reward that convinces the ruling elite (i.e. the *winning coalition*) to defect from the current leader (Besley and Kudamatsu 2008). Since the elite cannot be sure to be part of the winning coalition under a new ruler, and thus might loose the privileges enjoyed in the established system, supporting the incumbent might be the dominant strategy of the elite irrespective of the dictator's policy choices.

Casper and Tyson (2014) integrate public protests into a theory of elite coordination. In their model, popular uprisings act as a public signal that facilitates coordination among the elite through two channels: First, protests provide elite members with commonly observed information about regime legitimacy and the effectiveness of the leadership's coop-proofing strategies. Second, the sheer publicity of this signal improves the individual's belief about the actions of other elite members, thereby reducing uncertainty about the participation of compatriots in a coup plot.

In light of the selectorate approach, Apolte (2015) shows that a winning coalition that is trapped by loyalty can exploit a public rebellion for its own means. Whether or not the winning coalition keeps backing the incumbent government in case of a rebellion depends on the expected payoffs (Besley and Kudamatsu 2008). If coup leaders can gain public support and claim to enforce the people's will, the probability of being part of the winning coalition under a new leadership increases and therefore also the likelihood of maintaining private benefits. Hence, protests can provide an solution to the loyalty trap, given that there is a general interest of the winning coalition in overthrowing the incumbent.

In a similar model setup, Gilli and Li (2015) emphasize that a "revolutionary threat" (Acemoglu and Robinson 2006) can induce the ruling elite to replace the current leadership in order to avoid more drastic regime changes associated with a successful revolution. The underlying assumption in this model is that coups redistribute power within the ruling class, while mass revolts entail deeper political changes like democratization. Thus, regime change is the worst possible outcome for both, the dictator and the elite. If the incumbent dictator chooses inefficient economic policies that spur public grievances, the threat of a mass revolution will, under certain circumstances, induce the elite to mount a coup in order to calm down the public and avoid more dramatic regime changes. Wig and Rød (2014) expand this argument to election outcomes as a source of political instability in autoritarian regimes when incumbents reveal electoral weakness. Since such signals of weakened regime capacity are observed by the public, coups might be initiated by the elite to avoid public mass uprisings, either by serving as concessions to the opposition or by re-installing an effective repressive rule.

Yet, whether or not the signaling effect of social unrest on the alignment of elite expectations is sufficient to overcome the coordination obstacles of coup plotters depends on the type and strength of this signal. Johnson and Thyne (2016) argue that the signal from domestic actors in support of a coup is strongest when events of social unrest are located in urban areas and near the capital. Popular uprisings in larger cities are more likely to garner the attention of regime elites, especially when media freedom is restricted or media coverage is limited in peripheral areas (Casper and Tyson 2014). Also, protests in urban areas are usually directed more unambiguously against the incumbent government, while riots in rural areas often involve clashes between competing ethnic and tribal groups over resources (Almer, Laurent-Lucchetti, and Oechslin 2015). Since collective action is more easily organized in urban areas where urban dwellers are concentrated and the formation of crowds with coordinated beliefs is more likely than in sparsely populated zones, manifestations of popular discontent in urban areas impose a larger threat on regime elites because escalation of such protests into large-scale revolutions is relatively likely. Thus, elites (in particular military forces) cannot ignore popular mass movements in urban areas, but must act by either remaining loyal or by siding with the protesters (Johnson and Thyne 2016).

Overall, the theoretical literature consistently predicts that (urban) protests increase the likelihood of a coup attempt. In particular, in all of the presented arguments, public unrest acts as a proximate cause of actual coup attempts, apart from deeper structural causes like regime type, economic development and military capacity. In this sense, public protests open a temporally limited *window of opportunity* for regime elites by increasing both the ability and the disposition of elites to overthrow an incumbent leadership.

Related Empirical Literature

Fueled with new interest by the close interplay of public uprisings and military intervention in politics in the "Arab Spring" countries, the link between public protests and coup attempts has also been investigated in recent empirical studies. Applying extreme bounds analysis, Gassebner, Gutmann, and Voigt (2016) show that riots and political instability are robust predictors of coup risk. Examining more closely the transmission channels of the effect of social unrest on elite coordination and concerted action, Johnson and Thyne (2016) find that urban and non-violent protest events have a stronger effect on the likelihood of a coup attempt than peripheral and violent protests. In line with these findings, Casper and Tyson (2014) show that the signaling effect of public protests on the alignment of elite perceptions is most effective in regimes with better media freedom. Powell (2012) finds that popular revolts increase the likelihood that a coup attempt is successful.

Aside from political instability, Gassebner, Gutmann, and Voigt (2016) identify as most robust determinants of a coup attempt economic crises and the historical coup experience of a country, whereas the level of economic development, democratic institutions and military characteristics are not robustly related to coup risk. The importance of economic shocks for the disposition and ability of regime elites to stage a coup has also been pointed out by Londregan and Poole (1990) and Galetovic and Sanhueza (2000), but recent evidence is more ambiguous (Powell 2012). Interestingly, Kim (2014) shows that the effect of economic growth on the propensity of a coup attempt partly operates through public protests. There are also endogenous forces of instability inherent in different authoritarian regime types that facilitate the organization of a coup. For example, Geddes (1999) argues that military regimes are more vulnerable to internal policy differences or rivalries than personalist or one-party systems. While leaders in the latter two regimes have an incentive to cooperate with competing factions, officers in military regimes tend to split and replace concurrent ruling factions by coups (Svolik 2013).

Most of the existing studies do not explicitly take into account the potential endogeneity of coup attempts and public protests. (Anticipated) coup plots are likely to have feedback effects on citizens' perceptions about the regime's capacity to withhold policy concessions and repress opponents. Thus, the expectation of a coup attempt could itself lead to public revolts. Also, underlying factors that are not observed could drive both coups and protests. Those factors might be related to the overall legitimacy of the incumbent government as well as the effectiveness of repression. In the presence of omitted variables or reverse causality, ordinary least squares (OLS) as well as (conditional) logit estimates are biased and inconsistent. Two studies that discuss endogeneity problems associated with protests and coups are Casper and Tyson (2014) and Johnson and Thyne (2016). Casper and Tyson (2014) estimate a seemingly unrelated regression (SUR) model that controls for correlated error terms of the two simultaneously determined estimation equations on coups and protests. Johnson and Thyne (2016) use monthly data at the country level that allows for modeling more accurately the time structure of events. Yet, these approaches do not account for unobserved time-varying heterogeneity influencing both, coup attempts and public revolts, and thus do not sufficiently rule out endogeneity concerns.

Our paper thus contributes to the existing literature in that it addresses the problem of endogeneity by isolating exogenous variation in urban protests through deviations in rainfall in a particular set of countries, namely Sub-Saharan African countries. There is an increasing body of empirical literature on the relationship between climate and conflict[2]. In a seminal paper, Miguel, Satyanath and Sergenti (2004) use rainfall growth as an instrumental variable for estimating the effect of economic growth on civil wars in SSA. Similarly, weather-related instruments have been used in a variety of studies on the effects of adverse economic conditions on political change showing that income shocks can trigger democratization (Brueckner and Ciccone 2011; Burke and Leigh 2010).

The relationship between precipitation and lower-scale forms of social conflict like public protests and riots is analyzed in detail by Aidt and Leon (2015) and Hendrix and Salehyan (2012). Aidt and Leon (2015) study the effect of riots on democratic change, instrumenting riots by droughts. Their results indicate a strong correlation between extreme negative weather shocks and public protests which is particularly pronounced when protest events are located in urban areas. Similarly, Hendrix and Salehyan (2012) find that deviations in rainfall from long-run means increase the incidence protests. Both studies argue that public revolts are triggered by variations in rainfall levels through their effects on resource competition, food shortages and reduced state capacity. This argument is supported by Raleigh, Choi, and Kniveton (2015) who show that rainfall deviations increase social conflicts through increased food price volatility. Since the vast majority of food consumed in Africa is from domestic producers, food prices in Africa are generally locally determined (with few exceptions like rice) and thus, local climate change indeed affects food prices. While rainfall shocks impose economic hardship on net consumers of food in both, rural and urban areas, some producers and traders might also benefit from rising food prices. Yet, urban dwellers particularly suffer from food shortages if there is a lack of food supply from the periphery in the presence of a crop shortfall (van Weezel 2016). In fact, Smith (2014) finds that rising food prices lead to increased urban protests and Barrett (2013) shows that "food price" riots are overwhelmingly an urban phenomenon. Raleigh, Choi, and Kniveton (2015) argue that not all places are equally vulnerable to the adverse effects of anomalous climate events and the associated increased price volatility and show that these effects are particularly pronounced in market places that are usually located in urban areas. Since the response of food prices to anomalous weather shocks depends on food policies and might be thus endogenous to the incumbent political regime (Hendrix and Haggard 2015), we concentrate on exogenous variation in precipitation as the underlying cause of temporal hardship in urban areas.

Finally, the present paper is also related to the literature on conflicts and political change (Blattman and Miguel 2010) and in particular to the relationship of coup d'états

[2] For a general overview of this literature, see Burke, Hsiang and Miguel (2014).

and democratization. While the theoretical literature suggests that coup leaders have little incentives to extend political participation once in power, recent empirical literature shows that coups in autocratic settings actually do promote democratization under certain circumstances (Thyne and Powell 2014; Marinov and Goemans 2014; Powell 2012). Public protests provide a possible link to explain under which conditions coup attempts lead to democratic change. Following Gilli and Li (2015), mass movements can play an important role in shaping the incentives of coup leaders to implement (seemingly) democratic institutions. While elites might be able to exploit small-scale riots in order to seize and extend autocratic power, large-scale protests could force coup leader to make democratic concessions in order to gain legitimacy and avoid future coup-traps (Londregan and Poole 1990; Aidt and Leon 2015). Hence, understanding the relationship between public protests and coups also helps to explain autocratic regime stability and the underlying forces of political transition.

3 Data and Measurement

To test the relationship between protests and coups, we combine information on urban protests, coups, rainfall measures and economic factors for a sample of 39 Sub-Saharan African countries over the period from 1990 to 2007. We define our unit of analysis as the country-year. The main dependent variable *coup* is defined as a binary indicator of whether there is a coup attempt in a given country-year. We rely on a dataset provided by Powell and Thyne (2011) who define coup attempts as "illegal and overt attempts by the military or other elites within the state apparatus to unseat the sitting head of state using unconstitutional means" (p.252). The coding procedure of Powell and Thyne has several advantages to other datasets: First, it carefully differentiates coup attempts from other types of anti-government activities, such as international intervention, riots and public protests. Second, it includes not only successful, but also failed coup attempts. Third, it does not limit coup attempts to originate from the military but also considers coup attempts perpetrated by civilian members of the government. In addition, since the definition focuses on *illegal* activities, it differentiates coups from leadership changes resulting from (legal) political pressures, which might in turn be the consequence of public protests. Finally, the coding procedure does not apply a minimum death threshold, thereby also considering non-violent coup attempts where the threat of force suffices to induce the incumbent to step down. As an alternative measure we construct a count variable of all coup attempts in a given country-year (*coupcount*). For robustness checks, we also test alternative coup measures from the Center of Systemic Peace (CSP) Coup d'État Events Database (Marshall and Marshall 2015).

Data on public protests in SSA is available from the Social Conflict in Analysis Database (SCAD) version 3.1, updated on November 20, 2014 (Salehyan and Hendrix 2014). The data set contains information on different types of social conflict in Africa and their geo-referenced location from 1990 until 2013 and is based on Lexis-Nexis searches. To avoid conflation of coup attempts and public protest, we restrict our definition of public protests to the following categories of events: Organized/spontaneous demonstrations, organized/spontaneous riots, general/limited strike, anti-government violence and extra-government violence[3]. Importantly, these protest events do not include government forces as perpetrators and can be interpreted as a general signal of reduced regime legitimacy. Our protest variable captures the intensity of social unrest by measuring the logged sum of the duration (in days) of all relevant protest events in a given country-year. Since we concentrate on urban unrest, we include only protest events that are located in areas with a population density of more than 100 individuals per square kilometer. This weighted protest variable is taken from Aidt and Leon (2015) who employ GIS maps of African population distribution to determine whether a protest as defined above is located in a 1 square kilometer cell with more than 100 individuals.

To obtain the rainfall-related IVs, we use data from Miguel et al. (2004, 2011) and Ciccone (2011) that originate from the Global Precipitation Climatory Project (GPCP) for the period 1990 until 2009[4]. We compute our baseline instrument *drought* as a binary variable that equals one if average precipitation in a given country-year is below the 20th percentile (or 459.26 mm per year) of the sample distribution of annual rainfall. As can be seen from figure 1, drought years are distributed unevenly across SSA countries, with most droughts occurring in countries located in the Sahel and the Southern part of the continent. To account for this spatial heterogeneity of precipitation, we interact our drought instrument with region dummies for Western, Eastern, Middle and Southern Africa[5].

[3]Anti-government violence refers to violence against government targets by permanent or semi-permanent militias, whereas extra-government violence refers to violent events where government forces are neither an actor nor a target (for example, communal conflicts) (Salehyan et al. 2012).

[4]The GPCP combines rainfall data from gauge measures of weather stations with satellite information and provides monthly precipitation estimates at 2.5 latitude and longitude degree intervals. To ensure consistency with previous work, we draw on data from Ciccone (2011). In his dataset, precipitation is aggregated at the country-year level and measured as average precipitation in millimeters per year. Since Ciccone (2011) extends the dataset constructed by Miguel et al., applying their coding procedure, see Miguel et al. (2004) for details.

[5]The region classification is obtained from the United Nations Statistical Division. We have data for Sudan which is usually assigned to North Africa. Yet, since our analysis does not include other

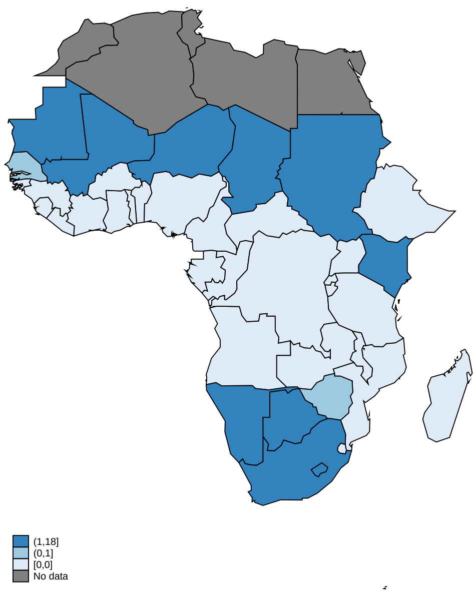


Figure 1: Total number of drought years by country in SSA

As most important control variable we include GDP per capita growth from the World Development Indicators (WDI) (World Bank 2015). Since economic growth is likely to be endogenous to coup risk and might impact on the link between riots and coups, we use variation in international commodity prices as instrument for economic growth. More precisely, we include annual changes in a country's extractive commodity price index as constructed by Bazzi and Blattman (2014). This index is a geometric average of international export prices of extractive commodities (such as oil, iron and gas) which is weighted by lagged country-specific export shares for each commodity. This procedure accounts for the fact that economies more dependent on extractive commodity exports are more sensitive to price shocks and ensures both, within-country variation through changes in international prices, and between-country variation through different trade shares.

In order to analyze systematic differences in the relationship between protests and coup risk across different political regime types, we use the combined Polity2-score from the Polity IV project (Marshall, Gurr, and Jaggers 2014) as well as the regime classification of Cheibub et al. (2010) that distinguishes between military, civil and royal autocracies. In some specifications, we include additional control variables. Data for (logged) GDP p.c. levels and total military expenditure is obtained from the WDI. In addition, we

North African countries, we assign Sudan to Middle Africa. For an overview of sample countries and regions, see table A.2.

calculate the time past since the last coup attempt (in years), its square and cubic term based on the data provided by Powell and Thyne (2011). Appendix tables A.1 and A.2 display sample countries and summary statistics.

4 Empirical Strategy

To investigate the link between public protests and coup attempts, we specify our structural specification as follows:

$$\begin{aligned} coup_{i,t} = \alpha + \beta ln(urban\ protest)_{i,t} + \gamma GDP\ p.c.\ growth_{i,t} \\ + \mathbf{X}'_{i,t}\Psi + \theta_i + \phi_t + \varepsilon_{i,t} \end{aligned} \tag{1}$$

where *coup* is a binary variable indicating whether there is a coup attempt in a given country-year and *ln(urban protest)* is defined as the logged duration (in days) of all protest events located in urban areas in a given country-year. Country (θ) and year (ϕ) fixed effects are included and the error term (ε) is assumed to be well behaved. In some specifications, additional control variables are included, captured by the vector $X_{i,t}$. These are GDP p.c. levels, the degree of democratic institutions reflected by the combined Polity2-value, military expenditure and a polynomial term of the time past since the last coup attempt. In the most rigid baseline specification, however, only economic growth is included as control variable to avoid distortion of the IV estimates by "bad controls" (Angrist and Pischke 2008; Burke, Hsiang, and Miguel 2014). Since coup attempts are expected to be influenced by recently observed changes in the environment of the elite, we regress coups on the contemporaneous protest intensity and economic growth.

The main problem associated with the estimation specification is that urban protests are likely to be endogenous to coup risk. (Anticipated) coup attempts might weaken popular perceptions about regime capacity, providing a signal for concerted action to be feasible (Miller 2012). Also, there might be unobserved factors affecting both, public protests and coup risk. To isolate exogenous variation, we use *drought* as an instrument for protests. Besides the technical advantages of this approach for causal inference, instrumenting public uprisings with (unexpected) weather shocks also has an appealing theoretical interpretation. Individual opponents of the regime face pronounced collective-action problems when coordinating concerted public action. These problems are similar to the coordination problems of regime elites, and might be even worse for the general public as has been pointed out by Tullock's "paradox of revolution" (1971). Since popular uprisings can be understood as a public good with substantial individual cost and general public benefits, rational individuals have little incentive to participate in such movements. Still, protests occur and often inhibit a stochastic moment. Threshold models of collective behavior suggest that public revolts are triggered by unexpected shocks that change the perceptions of a large part of the public regarding regime capacity and other individuals' actions, and this effect is particularly pronounced in urban areas (Kuran 1989; Apolte 2015). Weather-related shocks spur grievances especially in urban areas where people are concentrated and crowds with similar beliefs can be organized easily. Since negative rainfall shocks reduce agricultural output in non-irrigated systems (Dell, Jones, and Olken 2014), and urban populations strongly depend on domestic agricultural commodity supply, these shocks might thus trigger demonstrations, riots and revolutions.

While there are certainly other structural factors that influence the conflict potential of a society, isolating urban protest events that are triggered by weather-related shocks helps to identify variation in protests that is exogenous to coup risk. In general, for *drought* to satisfy the exclusion restriction as an IV, it must be uncorrelated with the error term of the structural equation. This means that droughts must have no effect on coup attempts through any channel other than public protests, *conditional* on control variables such as economic growth. Thus, negative weather shocks must not impact coup risk directly. In contrast to the vulnerability of large parts of the population to shocks in agricultural output in SSA countries, regime elites are assumed to rely on a more diversified income portfolio that makes them less vulnerable to weather shocks (Bazzi and Blattman 2014). In particular, elites usually control the rents from international trade in extractive resources and their income does thus not depend on agricultural commodities alone. For this reason, we expect that *drought* satisfies the exclusion restriction, which is supported by tests for overidentification shown below. For *drought* to be valid, the independence condition also requires that it is not driven by the dependent variable, i.e. coup attempts. Since weather phenomena are not determined by human activities, at least in the short run, this condition can be plausibly expected to be satisfied. Finally, for the empirical specification to be identified, droughts must be strongly correlated with public protest which is supported by evidence in Aidt and Leon (2015) and Hendrix and Salehyan (2012). The first stage regressions shown below confirm a high correlation between droughts and urban protests in our sample when we allow for regional heterogeneity in the effect of droughts on protests.

The inclusion of *GDP p.c. growth* is justified for two reasons: First, economic downturns reduce the opportunity cost of the elite to mount a coup and restrict the ability of the dictator to provide private benefits to his winning coalition. At the same time, GDP p.c. growth is likely to influence the occurrence of public protests as economic crises spur grievances among the population. Hence, economic growth is a potential confounder of the protest-coup nexus (Kim 2014). Second, negative weather shocks could influence the propensity of a coup risk also through the adverse effect on economic growth (Miguel, Satyanath, and Sergenti 2004). Thus, including GDP p.c. growth is necessary for *drought* to satisfy the exclusion restriction as instrument for public protest. Importantly, if *urban protest* remains statistically significant conditional on the inclusion of income growth, weather-driven urban protests affect coup risk through channels *other* than GDP per capita growth (e.g. through challenging state legitimacy). Yet, *GDP p.c. growth* is likely to be endogenous to coup risk as well. To account for this potential caveat, we employ extractive commodity price shocks from Bazzi and Blattman (2014) as instrument for economic growth. This identification strategy reflects the idea that regime elites are particularly sensitive to income fluctuations caused by changes in the rents from (extractive) resources (Kim 2014). Brückner et al. (2012) show that oil price shocks are strongly correlated with GDP growth and find positive and significant effects of oil-price-driven changes in GDP on political change and democratization. The exclusion restriction again requires that extractive commodity price shocks affect coup risk only through their impact on GDP per capita growth. In theory, changes in international prices could be affected by (anticipated) GDP growth or political instability (induced by coups) in large producer or consumer countries. To ensure that countries do not have price-setting power in world markets, Bazzi and Blattman (2014) exclude from a country's price index those commodities where the countries produces more than ten percent of global exports.

The first-stage equations for urban *protest* and *GDP p.c. growth* are then given by:

$$protest_{i,t} = \mathbf{Z}'_{i,t}\Lambda + \mathbf{X}'_{i,t}\Pi + \theta_i + \phi_t + \upsilon_{i,t} \tag{2}$$

and

$$GDP\ p.c.\ growth_{i,t} = \mathbf{Z}'_{i,t}\Theta + \mathbf{X}'_{i,t}\Upsilon + \theta_i + \phi_t + \nu_{i,t} \tag{3}$$

where $\mathbf{Z}_{i,t}$ is the vector of instruments that are excluded from the second stage including *drought*, its interaction with region dummies, and the *extractive price shock*. $\mathbf{X}_{i,t,}$ is a vector of controls that includes GDP p.c. growth (when not instrumented) and further additional covariates in some of the regressions.

5 Main Results

OLS and Reduced Form Estimates

Table 1 presents as benchmark results the ordinary least squares (OLS) and conditional logit estimates of our structural equation (1). The dependent variable in column (1) is our binary coup indicator. The point estimate of *urban protest* is positive and significant, indicating that a one-unit increase in the log of *urban protest* increases the probability of a coup attempt by 1.97 percentage points. At the sample mean of *coup*, which is 5.6 percent, this corresponds to a substantial relative increase in the propensity of a coup attempt of roughly 35 percent. The coefficient of *urban protest* is similar in column (2), where the dependent variable *coupcount* is a count measure of total coup attempts in a given country-year. The coefficient of *GDP p.c. growth* has the expected negative sign and is significant in all regressions, supporting the notion that economic growth tends to reduce the likelihood of a coup attempt (Kim 2014; Gassebner, Gutmann, and Voigt 2016).

Since coup is a binary indicator, we also estimate a fixed effects conditional logit regression in column (3). The estimates are similar to the OLS estimator with regard to signs and significance levels, but the magnitude of the coefficients is much larger and the sample size substantially reduced. Even though the logit estimator captures more adequately the binary nature of our dependent variable, we continue with least-squares estimation techniques because instrumenting endogenous regressors is not feasible in logit models. Angrist (2001) and Angrist and Pischke (2008) show that OLS often perform well even for limited dependent variables, which is supported by the qualitatively similar estimates in table 1.

In column (4), we include lagged values of *urban protest* to control for a potentially delayed effect of public unrest on coup risk. Yet, the insignificant coefficient suggests that protests have an immediate rather than a postponed impact on the probability of a coup attempt. This is in line with the idea that protests open a (temporally limited) window of opportunity for changing power positions within the ruling elite.

Table 1: Structural and Reduced Form Regressions

	Structural Regressions					Reduced Form Regressions	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dependent variable:	Coup_t	Coupcount_t	Coup_t	Coup_t	Coup_t	Coup_t	Coupcount_t
	OLS	OLS	Logit	OLS	OLS	OLS	OLS
Urban protest_t, logs	0.0197**	0.0205*	0.477**	0.0269**	0.0191**		
	(0.00892)	(0.0103)	(0.222)	(0.0108)	(0.00874)		
Urban protest_{t-1}, logs				-0.0114			
				(0.00918)			
GDP p.c._t, logs					-0.0557		
					(0.0819)		
Polity2_t					-0.0107***		
					(0.00352)		
Military expenditure_t, logs					0.00143		
					(0.00507)		
Drought_t						0.0589***	0.0734***
						(0.0133)	(0.0164)
$\text{Drought}_t \times$ Eastern Africa						-0.0676**	-0.0703**
						(0.0309)	(0.0332)
$\text{Drought}_t \times$ Middle Africa						-0.0679***	-0.106***
						(0.0220)	(0.0324)
$\text{Drought}_t \times$ Southern Africa						-0.0567**	-0.0617**
						(0.0240)	(0.0287)
GDP p.c. growth_t	-0.795***	-0.906***	-11.17***	-0.917***	-0.754***	-0.803***	-0.915***
	(0.268)	(0.327)	(3.603)	(0.264)	(0.259)	(0.270)	(0.327)
Country FE	Y	Y	Y	Y	Y	Y	Y
Time FE	Y	Y	Y	Y	Y	Y	Y
Time since last coup	N	N	N	N	Y	N	N
Time since last coup^2	N	N	N	N	Y	N	N
Time since last coup^3	N	N	N	N	Y	N	N
(Pseudo) R-squared	0.056	0.049	0.173	0.071	0.108	0.052	0.047
Observations	682	682	324	643	585	682	682

Note: Robust standard errors reported in parentheses. OLS = Ordinary least square estimations. Column (3) reports conditional logit coefficients. Significance levels: *** p<0.01, ** p<0.05, * p<0.1.

Finally, in column (5) we include additional control variables that have been discussed to be robust predictors of coup attempts in the literature (see for example Gassebner, Gutmann, and Voigt (2016)). As expected, higher GDP p.c. levels lower the likelihood of a coup attempt, though the coefficient is not significant. More democratic institutions – reflected by higher Polity2-values – significantly reduce the propensity of a coup, which is in line with the observation that coup attempts are less frequent in democracies. In contrast, higher levels of military expenditure are positively associated with coup risk, although the coefficient again is not significant. We also include a polynomial term for the time past since the last coup attempt (coefficients not shown) and find that recent experience with coups increases the vulnerability of regimes which in line with evidence presented in Gassebner, Gutmann, and Voigt (2016).

However, the results in table 1 should not be interpreted as causal effects because endogeneity is a serious concern for both, *urban protest* and *GDP p.c. growth*. Thus, the OLS and logit estimates are likely to be biased. Before discussing the IV estimates, columns (6) and (7) of table 1 present reduced-form estimates on the link between our protest instruments, i.e. *drought* and its interaction with region dummies, and coup risk. The coefficients are highly significant both for the binary coup variable and the count indicator of coups, but show substantial heterogeneity across regions which will be discussed in more detail below.

2SLS Estimates

Table 2 presents the two-stage least-squares estimates (2SLS). The dependent variables are *coup* and *coupcount*. In column (3) and (6), we include additional control variables in our specification. Panel A reports the coefficients of the structural equation estimated by 2SLS. Panel B presents the first-stage results for *urban protest*, while Panel C presents the first-stage results for *GDP p.c. growth* (when relevant).

In column (1) to (3), drought and the interaction terms of drought and three region dummies (Eastern Africa, Middle Africa, Southern Africa) are introduced as instrument for *urban protest*.[6] The first-stage results in panel B, column (1) show that droughts are significantly associated with the incidence of urban protests, but that there is substantial heterogeneity across regions. In Western Africa (which is the omitted baseline region), a drought increases the average duration of protest events in a given country-year by roughly 14 days.[7] In other parts of Sub-Saharan Africa, this effect is much less pronounced, yet droughts always lead to a significant positive increase in the duration of urban protests. The F-statistic on the excluded instruments indicates that negative weather shocks are strong predictors of social unrest in urban areas.

The observed regional heterogeneity stems from the fact that extreme weather shocks are distributed unevenly across Sub-Saharan Africa with droughts being much more common in Western Africa (8.1 percent of sample years) and to a lesser extend in Southern, Middle and Eastern African countries (4.8, 3.7 and 0.4 percent of sample years re-

[6] Western Africa is the omitted category. For an overview of sample countries included in each regional category, see appendix table A.2.

[7] Since we include the natural logarithm of *urban protest*, a drought increases the duration of urban protests by $exp(2.642) = 14.041$ days.

spectively). Even though SSA countries in general suffer from low precipitation levels, countries less prone to extreme negative rainfall shocks may be better able to compensate for this infrequent phenomenon with improved water storage, irrigation systems and thus less dependence of agricultural output on rainfall. Moreover, we hypothesized that negative weather shocks trigger urban protests mainly through foot shortages and rising food prices (Raleigh, Choi, and Kniveton 2015). Yet, the reaction of food prices to droughts might vary across countries and regions and depends on food policies, trade openness and a country's dependence on food imports (Hendrix and Haggard 2015). Figure A.1 shows that the correlation between total drought years and food prices indeed varies across regions. While this correlation is positive for Western and Southern African countries, there is a negative correlation in Middle and Eastern African countries. Thus, a higher number of extremely dry years leads to higher food price volatility in Western and Southern Africa which in turn might spur urban grievances, whereas we find no such effect in Middle and Eastern Africa. The food price data shown refers to the country-specific average of monthly deviations from a food price index provided by van Weezel (2016) that does not vary over time, and thus serves only as a broad proxy for the response of food prices to weather shocks. Yet, the descriptive pattern points to a heterogeneous relationship between droughts and food prices that in turn helps to explain the regional heterogeneity displayed in the first stage results for *urban protest*.

Turning to our two-stage least-squares results for the effect of urban protests on coup attempts in panel A of column (1), we find that weather-driven urban protests significantly increase the likelihood that a coup is attempted. Again, the coefficient of economic growth is negative and significant. The estimates closely resemble the OLS results from table 1. In column (2) of table 2 we use *coupcount* as our dependent variable, and in column (3) we include additional control variables as in model (5) of table 1. The point estimates for *urban protest* are slightly larger than the OLS estimates and significant at the five percent level. Yet, so far we have not controlled for potential endogeneity of GDP p.c. growth. Related work suggests a link between rainfall, economic growth and conflict (Miguel, Satyanath, and Sergenti 2004; Kim 2014). Thus, the coefficients of *urban protest* might in part be driven by the impact of rainfall deviations on (transitory) income shocks. To account for the potential endogeneity of per capita income growth, we include *extractive price shock* as additional instrument in our regressions.

The results for our baseline specifications are presented in columns (4) to (6) of table 2. Regarding the first-stage results for *GDP p.c. growth* that are reported in panel C, extractive commodity price shocks have a negative and significant impact on economic

Table 2: 2SLS, Public Protests and Coups

	(1) Coup$_t$	(2) Coupcount$_t$	(3) Coup$_t$	(4) Coup$_t$	(5) Coupcount$_t$	(6) Coup$_t$
			A. 2SLS			
Urban protest$_t$, logs	0.0190*	0.0275**	0.0274**	0.0207*	0.0293**	0.0316**
	(0.00997)	(0.0136)	(0.0129)	(0.0110)	(0.0147)	(0.0145)
GDP p.c.$_t$, logs	-0.795***	-0.903***	-0.753***	-0.214	-0.192	0.326
	(0.244)	(0.279)	(0.245)	(1.192)	(1.392)	(1.209)
Weak IV-Test: rel. bias >10%						
Kleibergen-Paap F statistic	28.28	28.28	21.14	1.412	1.412	0.935
Stock-Yogo critical value	10.27	10.27	10.27	8.78	8.78	8.78
A-R Wald, F (p value)				[0.0215]	[0.0500]	[0.0736]
A-R Wald, χ^2 (p value)				[0.0170]	[0.0417]	[0.0582]
Hansen overid. test (p value)	0.931	0.864	0.574	0.975	0.880	0.569
Additional controls	N	N	Y	N	N	Y
Country & year FE	Y	Y	Y	Y	Y	Y
Observations	682	682	585	682	682	585
			B. First stage for *urban protest$_t$*			
Drought$_t$	2.642***	2.642***	2.557***	2.651***	2.651***	2.570***
	(0.251)	(0.251)	(0.287)	(0.246)	(0.246)	(0.280)
Drought$_t\times$ Eastern Africa	-2.306***	-2.306***	-2.253***	-2.299***	-2.299***	-2.250***
	(0.579)	(0.579)	(0.643)	(0.579)	(0.579)	(0.629)
Drought$_t\times$ Middle Africa	-2.601***	-2.601***	-2.583***	-2.610***	-2.610***	-2.616***
	(0.266)	(0.266)	(0.286)	(0.260)	(0.260)	(0.277)
Drought$_t\times$ Southern Africa	-2.387***	-2.387***	-2.401***	-2.404***	-2.404***	-2.419***
	(0.328)	(0.328)	(0.354)	(0.324)	(0.324)	(0.348)
Extractive price shock$_t$	–	–	–	-0.021	-0.021	-0.0518
				(0.0250)	(0.0250)	(0.0355)
GDP p.c. growth$_t$	Y	Y	Y	–	–	–
Additional controls	N	N	Y	N	N	Y
Country & year FE	Y	Y	Y	Y	Y	Y
F-statistic on excl. IVs	28.28	28.28	21.14	27.64	27.64	22.29
			C. First stage for *GDP p.c. growth$_t$*			
Drought$_t$				-0.002	-0.002	0.000154
				(0.00757)	(0.00757)	(0.00980)
Drought$_t\times$ Eastern Africa				-0.0236	-0.0236	-0.0203
				(0.0267)	(0.0267)	(0.0274)
Drought$_t\times$ Middle Africa				-0.00789	-0.00789	0.00930
				(0.0206)	(0.0206)	(0.0224)
Drought$_t\times$ Southern Africa				0.00859	0.00859	0.00546
				(0.0116)	(0.0116)	(0.0137)
Extractive price shock$_t$				-0.00507**	-0.00507**	-0.00458*
				(0.00230)	(0.00230)	(0.00245)
Additional controls				N	N	Y
Country & year FE				Y	Y	Y
F-statistic on excl. IVs				1.47	1.47	0.98

Note: 2SLS = Two-stage least-squares. In panel A, weak-instrument robust inference is reported for two significance tests of the endogenous regressors in the structural regression (p values in square brackets). The respective test statistics are Anderson-Rubin's Wald F and Wald χ^2 statistic. The null hypothesis is that the endogenous regressors jointly equal zero and that the overidentifying restrictions (where relevant) are valid. In addition, Hansen's test of overidentifying restrictions is reported (p value). The null is that the instruments are valid and correctly excluded from the structural regression. Hansen's J statistic is robust to heteroskedasticity and autocorrelation. Robust standard errors are reported in parentheses (). The omitted regional baseline category in panel B and C is *Western Africa*. The additional control variables in column (3) and (6) are GDP p.c. (logs), Polity2, military expediture (logs) and a polynomial term for the time past since the last coup attempt. Significance levels: *** p<0.01, ** p<0.05, * p<0.1

growth. This result is somewhat troubling as it implies that positive changes in export-weighted extracive commodity prices reduce economic growth rates. While empirical findings of a "resource curse" in principle provide an explanation for a negative relationship between resource rents and economic development, the focus is usually on the long-run link between resources and growth (Mehlum, Moene, and Torvik 2006; Collier and Goderis 2012). Yet, the negative correlation found in our sample is in line with comparable evidence in Aidt and Leon (2015).

Importantly, droughts do not explain per capita income growth in this sample which suggests that the exclusion condition for our instruments is valid, i.e. that adverse weather shocks influence coup attempts only through public protests. However, the Kleibergen-Paap F statistic in column (4) to (6) is below the critical value for avoiding weak instruments as suggested by Stock and Yogo (2002). Thus, the included instruments might be weak. From the F-statistics on the excluded instruments in the first stage regressions in panel B and C we conclude that concerns of weak IVs apply mainly to GDP p.c. growth. In the second stage regression, we therefore also report weak-IV robust inference to assess significance. The null hypothesis of the Anderson-Rubin Wald tests is that the endogenous regressors jointly equal zero and that the overidentification restrictions are valid. The corresponding p values (in square brackets) in panel A indicate that the coefficients of *urban protest* and *GDP p.c. growth* are jointly significant at the 5% or 10% level.

The coefficients of *urban protest* in column (4) to (6) of table 2 are somewhat larger than the OLS estimates suggesting that part of the effect of rainfall-driven economic crises is indeed channeled through public protests as found by Kim (2014). Taking the point estimates in column (4), a drought occurring in a Western African country leads to an increase in the log of *urban protest* of 2.651, while a one-unit increase in the log of *urban protest* raises the probability of a coup attempt by 0.0207. Thus, a drought in Western Africa on average leads to an increase in the likelihood of a coup attempt of $0.0207 \times 2.651 = 0.0549$ or 5.49 percentage points (through the increased duration of urban protests). Evaluated at the sample mean of coup attempts, this effect corresponds to a doubling of coup risk during drought years.

Since five IVs are included for two endogenous variables, the overidentifying restrictions can be exploited to test whether the exclusion condition holds, i.e. whether the instruments are uncorrelated with the error term of the structural regression. The reported test statistic is Hansen's J that is consistent in the presence of heteroskedasticity and autocorrelation. For all specifications in table 2 the test fails to reject the null hypothesis that the excluded instruments are valid and correctly excluded from the structural regression.

6 Robustness

Alternative Protest Measures and Instruments

To assess the robustness of the 2SLS results, we include alternative protest indicators in columns (1) to (3) of table 3. The dependent variable in table 3 is the binary coup indicator. We start by comparing the effect of urban versus rural protest events. To this end, we construct the variable *rural protest* that is the sum of protest events located in areas with less than 100 individuals per km^2 and include it in our baseline specification in addition to *urban protest.* Thus, the baseline category in column (1) refers to country-years without any protest event. We find support for our expectation that urban protests have a stronger effect on coup attempts as compared to protests that are located in peripheral areas. The coefficient on *urban protest* is more than three times greater than the coefficient on *rural protest*, and both effects are significant when applying weak-IV robust inference as reported by the Anderson-Rubin p values. In column (2), we include the variable *protest, not weighted* that is the logged sum of all protest events, irrespective of the location of these events. The results are robust to our baseline estimates. In column (3), we include an alternative indicator of urban protests that weights events according to the distance of their location to the capital and is obtained from Aidt and Leon (2015). As argued by Johnson and Thyne (2016), the signaling effect of public unrest on elite perceptions should be stronger when the protests are centered near the capital. Indeed, the coefficient of *protest, near capital* is positive and large. Yet, the first-stage results for all three alternative protest variables display small F-statistics on the excluded IVs and thus indicate that droughts are more suitable to predict the incidence of *urban* protests than other forms of public unrest.[8]

In columns (4) and (5), we test alternative rainfall-related IVs. Since precipitation is expected to influence urban unrest through losses in agricultural output and associated food shortages, the impact of droughts should be stronger in countries that are more dependent on agriculture. Following Burke and Leigh (2010), we therefore interact our drought indicator with the country-specific share of agricultural employment in total employment at the beginning of the sample period (i.e. in 1990). Data on agricultural

[8]The first-stage results for *rural protest* and *GDP p.c. growth* are reported in appendix table A.3.

Table 3: 2SLS, Alternative Protest Variables & Instruments

	(1) Coup$_t$	(2) Coup$_t$	(3) Coup$_t$	(4) Coup$_t$	(5) Coup$_t$
			A. 2SLS		
Urban protest$_t$, logs	0.0185			0.0674	0.0302**
	(0.0211)			(0.0464)	(0.0137)
Rural protest$_t$, logs	0.00529				
	(0.0350)				
Protest$_t$, not weighted (logs)		0.0215			
		(0.0164)			
Protest$_t$, near capital (logs)			0.0395		
			(0.0447)		
GDP p.c. growth$_t$	-0.316	-0.524	-0.483	0.557	0.445
	(1.644)	(1.220)	(1.377)	(0.972)	(0.837)
Weak IV-Test: rel. bias >10%					
Kleibergen-Paap F statistic	0.254	1.889	2.100	1.088	1.299
Stock-Yogo critical value	6.61	8.78	8.78	7.03	7.56
A-R Wald, F (p value)	[0.0215]	[0.0215]	[0.0215]	[0.110]	[0.0324]
A-R Wald, χ^2 (p value)	[0.0170]	[0.0170]	[0.0170]	[0.0999]	[0.0256]
Hansen overid. test (p value)	0.923	0.886	0.853	–	0.414
Country & year FE	Y	Y	Y	Y	Y
Observations	682	682	682	538	538
			B. First stage for protest variable$_t$		
	Urban protest$_t$	Protest$_t$, not weighted	Protest$_t$, near capital	Urban protest$_t$	
Drought$_t$	2.651***	1.802***	0.738***		
	(0.246)	(0.364)	(0.171)		
Drought$_t \times$ Eastern Africa	-2.299***	-1.820**	-0.908*		
	(0.579)	(0.826)	(0.502)		
Drought$_t \times$ Middle Africa	-2.610***	-2.626***	-1.182***		
	(0.260)	(0.630)	(0.357)		
Drought$_t \times$ Southern Africa	-2.404***	-1.559***	-0.608**		
	(0.324)	(0.522)	(0.264)		
Drought$_t \times$ agricultural employment				0.0121	0.0650***
				(0.00828)	(0.00665)
Drought$_t \times$ agri. emp.$\times$ Eastern Africa					-0.0589***
					(0.0107)
Drought$_t \times$ agri. emp.$\times$ Southern Africa					-0.0587***
					(0.00934)
Extractive price shock$_t$	-0.0210	-0.0840	-0.0345	-0.0245	-0.0258
	(0.0250)	(0.0548)	(0.287)	(0.238)	(0.0238)
Country & year FE	Y	Y	Y	Y	Y
F-statistic on excl. IVs	27.64	5.72	4.40	1.79	30.29

Note: 2SLS = Two-stage least-squares. Robust standard errors in parentheses (). In panel A, weak-instrument robust inference and Hansen's J statistic (p values) are reported as in table 2. ***Rural protest, logs*** is the logged duration of protest events located in areas with less than 100 individuals per km^2. ***Protest, not weighted (logs)*** is the logged duration of all protest events. ***Protest, near capital (logs)*** weights each protest with the inverse of the log distance between the location of the event and the capital city of the respective country (variable obtained from Aidt and Leon (2015)). ***Agricultural employment*** is the share of agricultural employment in total employment in 1990. The first stage results for ***rural protest, logs*** and ***GDP p.c. growth*** are reported in appendix table A.3. Significance levels: *** p<0.01, ** p<0.05, * p<0.1

Table 4: 2SLS, Alternative Measures of Coups and Alternative Sample Criteria

	(1)	(2)	(3)	(4)	(5)	(6)
	CSP coup data			PT coup data		
Dependent Variable	Coup_t	Successful coup_t	Failed coup_t	Successful coup_t	Failed coup_t	Coup_t, no civil war periods
Urban protest_t	0.0393**	0.00700	0.0323**	0.00600	0.0147	0.0126
	(0.0159)	(0.00774)	(0.0153)	(0.00811)	(0.00935)	(0.0104)
GDP p.c. growth_t	0.449	0.115	0.334	-0.132	-0.0817	-0.530
	(1.289)	(0.763)	(1.042)	(0.676)	(1.042)	(2.574)
Weak IV-Test: rel. bias >10%						
Kleibergen-Paap F statistic	1.412	1.412	1.412	1.412	1.412	0.430
Stock-Yogo critical value	8.78	8.78	8.78	8.78	8.78	7.56
A-R Wald, F (p value)	[0.00486]	[0.246]	[0.0379]	[0.223]	[0.111]	[0.105]
A-R Wald, χ^2 (p value)	[0.00348]	[0.226]	[0.0310]	[0.204]	[0.0972]	[0.0903]
Hansen overid. test (p value)	0.489	0.563	0.556	0.506	0.460	0.725
Country & year FE	Y	Y	Y	Y	Y	Y
Observations	682	682	682	682	682	551

Note: 2SLS = Two-stage least-squares. Robust standard errors in parentheses (). In panel A, weak-instrument robust inference and Hansen's J statistic (p values) are reported as in table 2. In column (1) to (3), data on coups is drawn from the Center of Systemic Peace (CSP), otherwise coup data comes from Powell and Thyne (2011). The instruments included in the first stage are *drought* and its interaction with dummy variables for *Eastern Africa*, *Middle Africa* and *Southern Africa* as well as *extractive price shock*. The first-stage results for *urban protest, logs* and *GDP p.c. growth* are reported in appendix table A.4. Significance levels: *** p<0.01, ** p<0.05, * p<0.1

employment is taken from the WDI. In panel B, column (4), we report the first-stage results for the effect of *droughts* × *agricultural employment*, while in column (5) we again interact this instrument with region dummies.[9] The first-stage regressions produce strong and significant coefficients when allowing for regional heterogeneity. The coefficient of *urban protest* in panel A of column (5) is significant at the five percent level and increases in magnitude compared to our baseline results. In contrast, the point estimate of *GDP p.c. growth* in the structural estimation changes from negative to positive.

Alternative Measures of Coups Attempts

To test whether our results are driven by a specific definition of coup attempts, we test different definitions and data sources of coup incidents in table 4. In columns (1) to (3), the measure of coup attempts is obtained from the CSP Coup d'Ètat Events database (Marshall and Marshall 2015). The point estimate of *urban protest* in column (1) is robust to the use of this different dataset. In column (2) and (3), we separately test the

[9]Due to a reduction in sample size when including data on agricultural employment, we have not enough observations to include an interaction term with Middle Africa.

effect of urban protests on successful and failed coup attempts. Powell and Thyne (2011) define a coup attempt as successful if "the perpetrators seize and hold power for at least seven days" (p.252). A similar definition is applied in the CSP dataset. Interestingly, we find a positive and significant effect of *urban protest* only for failed coup attempts, but not for successful coups. The same conclusion applies when we use data from Powell and Thyne (the point estimate of *urban protest* for failed coups is borderline significant (p value = 0.117)). Assessing the the coefficients based on weak-IV robust significance tests yield support to the notion that urban protests increase failed coup attempts but not successful coups. This result is interesting as it contradicts existing studies arguing that political instability and protests improve the chances of plotters to successfully stage a coup (Powell 2012). It suggests that the signal sent by urban unrest changes the disposition of elites for a coup, but not necessarily the ability to successfully enforce a leadership change. Popular uprisings could induce elites to overstate the vulnerability of the incumbent regime. Also, the public might not necessarily react favorably to coup plotters despite widespread grievances with the incumbent regime. The recent coup attempt in Turkey may serve as an example for such a situation. Overall, the results suggest that future research should distinguish more clearly the factors that increase the disposition of elite members for a coup conspiracy and the ability of coup plotters to effectively seize power.

In addition, the occurrence of public protests and coup incidents is likely to be blurred in periods of enduring civil war, where different types of violent conflicts tend to overlap and cannot be disentangled meaningfully. Therefore, periods of civil war as defined by the Armed Conflict Database (ACD) (Gleditsch et al. 2002) are excluded from the analysis in column (6) of table 4. This reduces the sample to 551 country-years, but the second-stage results again support the hypothesis that public uprisings trigger coup attempts.

The Role of the Political Regime

The regressions presented so far do not differentiate between political regimes. Consequently, the Polity2-scores of the country-years included in the sample range from -10 (Swaziland) to +9 (South Africa) and thus reflect the full range of political systems from repressive autocracies to relatively stable democracies. Even though coup attempts could, in theory, happen in democracies for the same hypothesized reasons as in autocracies, this form of political conflict is more widespread in non-democratic regimes that lack alternative channels of expressing political grievances like elections. Also, the risk of a coup attempt might differ across types of autocracies. Geddes (1999) points out that military dictatorships inhibit more endogenous sources of instability than other types of autocratic regimes. Thus, the effect of political uprisings in these regimes could have a larger impact on coup attempts than in personalist and one-party regimes or monarchies.

In table 5 we analyze whether the link between public protests and coup attempts differs systematically across political regimes. To this end, we present OLS and 2SLS results for two different specifications. We start by investigating the differential effect of urban unrest in democratic and autocratic systems. To this end, we construct the variable *initial Polity2 value* as a time-invariant indicator of the combined Polity2-score in 1990 and interact it with our protest variable. The negative point estimate of the interaction term in column (1) indicates that urban protests in countries with more democratic institutions (i.e. positive initial Polity2-values) have a significant negative effect on the risk of a coup attempt, whereas countries with autocratic institutions (i.e. negative Polity2-values) are associated with a positive link between protests and coups. This result is in line with existing studies (Wig and Rød 2014) and suggests that democracies provide different institutional channels that translate public grievances into political change (such as general elections) than non-democratic systems.

In column (2), we take a closer look at different types of non-democratic systems. To distinguish democracies from autocracies, Cheibub, Gandhi, and Vreeland (2010) define democracies according to the presence of free and contested elections and thus apply a more parsimonious definition of autocracy than (Marshall, Gurr, and Jaggers 2014). According to their categorization, all countries in our sample are characterized as autocracies in 1990. Moreover, the authors classify different types of dictatorship according to the nature and size of the ruling elite that backs the incumbent government and the major method of leadership removal. Accordingly, they distinguish between military regimes that are confronted with key potential rivals from the armed forces within juntas, royal regimes that rely on family networks, and civil regimes that create smaller bodies within a regime party to co-opt rivals.

It follows from this distinction that different types of dictatorships provide different incentives and constraints in the relationship between regime elites and the government. While military regimes tend to split when faced with unexpected problems, civilian regimes try to co-opt their critics (Geddes 1999). Yet, if both, the ability and disposition of elites to stage a coup differ across different types of authoritarian systems, the effect of urban grievances on ruling elites is also likely to vary across regimes. To test this hypothesis, we include the interaction terms of *urban protest* with two dummy variables

Table 5: Political Regimes, Urban Protests and Coups

	OLS		2SLS	
Dep. var.: Coup$_t$	(1)	(2)	(3)	(4)
Urban protest$_t$	0.000312	0.0268*	0.0118	0.0968
	(0.00606)	(0.0134)	(0.0144)	(0.372)
Urban protest$_t \times$ initial Polity2 value	-0.00317*		-0.00574	
	(0.00168)		(0.0130)	
Urban protest$_t \times$ initial civil regime		-0.0178		-0.0687
		(0.0145)		(0.353)
Urban protest$_t \times$ initial royal regime		-0.0566*		-2.714
		(0.0317)		(7.850)
GDP p.c. growth$_t$	-0.793***	-0.791***	-0.781***	-0.835***
	(0.268)	(0.270)	(0.236)	(0.242)
Weak IV-Test: rel. bias >10%				
Kleibergen-Paap F statistic			0.611	0.224
Stock-Yogo CV 10% max IV rel bias			8.78	n.a.
A-R Wald, F (p value)			[0.0414]	[0.0269]
A-R Wald, χ^2 (p value)			[0.0338]	[0.0220]
Hansen overid. test (p value)			0.768	0.724
Country & year FE	Y	Y	Y	Y
Observations	682	682	682	682

Note: OLS = Ordinary least squares. 2SLS = Two-stage least-squares. Robust standard errors in parentheses (). Weak-instrument robust inference and Hansen's J statistic (p values) are reported as in table 2. The 2SLS regressions in column (3) and (4) include as IVs ***drought*** and its interaction with the region dummies. Significance levels: *** p<0.01, ** p<0.05, * p<0.1

that indicate whether a political regime is characterized as a civil or royal autocracy in 1990. The main coefficient of *urban protest* reports the effect of social unrest in initial military dictatorships.[10] We find a positive effect of *urban protest* on the likelihood of a coup attempt in military regimes, but no significant link in civilian dictatorships, whereas the association between protests and coups is negative and significant in monarchies. Hence, in military regimes, where the government suffers from internal rivalries, competing factions seem to be both more capable and willing to exploit public grievances for their own means to stage a coup (Svolik 2013). In contrast, in civilian systems, regime parties serve as device to penetrate and control the society and thus make it more difficult for rivaling factions to defect as a reaction to public unrest. Finally, in monarchies the main threat for rulers originates from family members (Cheibub, Gandhi, and Vreeland 2010). Public grievances against the monarch usually extend to challenging the legitimacy of the whole royal family. In these regimes, urban unrest is likely to increase the cohesion among the ruling family network in order to avoid a revolution that could throw the whole royal family out of office. The 2SLS results presented in column (3)

[10]While the political regime can clearly change in the aftermath of a coups and thus is likely endogenous to coup risk, we expect that current coup attempts have no effect *ex-post* on the initial regime type.

and (4) broadly support our findings, though the coefficients are only significant when assessed with weak-IV robust inference.

7 Limitations and Conclusion

This paper analyzes the impact of urban unrest on the likelihood of coups d'ètat by isolating exogenous variation in protests through rainfall shocks. The presented evidence provides new and extended insights to the literature on the driving forces of coups and contributes to a better understanding of the link between public protests, political instability and coup attempts. Comparable empirical evidence supports the findings of this paper, though existing studies do not take into account the potential endogeneity of riots and coups (Casper and Tyson 2014; Powell 2012; Johnson and Thyne 2016).

Yet, there are some potential caveats of the empirical approach. Most importantly, the IVs chosen must be reliable to capture the causal link between public protest and coups. Since public protests and rainfall shocks are short-lived and local events (Almer, Laurent-Lucchetti, and Oechslin 2015), future research should focus on geographically and temporally disaggregated data to take into account timing and spatial considerations of the link between rainfall shocks and protests seriously (Raleigh, Linke, and Havard 2010; Harari and La Ferrara 2014). Geo-coded data on the exact location and timing of protest events should be combined with respective data on precipitation levels, since both might vary substantially within countries and over the course of a calendar year. Similarly, disentangling the time structure of protest events and coup attempts could help to establish the causal link between these two sources of political instability (Johnson and Thyne 2016). Even though our results offer only limited conclusions of the relationship between protests and coups due to the high level of aggregation, the findings point to an important causal mechanism between social unrest and elite perceptions.

As an extension of the analysis, it would be interesting to study the links between coups, public protest and democratization. Recent empirical findings show that coups in autocracies promote democratization under certain circumstances (Marinov and Goemans 2014; Thyne and Powell 2014). Public protests could play an important role in shaping the incentives of coup leaders to implement democratic institutions. While elites might be able to exploit small-scale riots in order to extend autocratic power, large-scale movements could force coup leaders to make democratic concessions in order to gain legitimacy and avoid future unrest.

References

Acemoglu, Daron, and James A. Robinson. 2006. *Economic Origins of Dictatorship and Democracy.* Cambridge University Press.

Aidt, Toke S, and Gabriel Leon. 2015. "The Democratic Window of Opportunity Evidence from Riots in Sub-Saharan Africa." *Journal of Conflict Resolution* Doi: 10.1177/0022002714564014.

Almer, Christian, Jeremy Laurent-Lucchetti, and Manuel Oechslin. 2015. *Water Scarcity and Rioting: Disaggregated Evidence from Sub-Saharan Africa.* MPRA Paper No. 34487.

Angrist, Joshua D. 2001. "Estimation of Limited Dependent Variable Models with Dummy Endogenous Regressors: Simple Strategies for Empirical Practice." *Journal of Business & Economic Statistics* 19 (1).

Angrist, Joshua D, and Joern-Steffen Pischke. 2008. *Mostly Harmless Econometrics: An Empiricist's Companion.* Princeton University Press.

Apolte, Thomas. 2015. *Autocracy and the Public: Mass Revolts, Winning Coalitions, and Policy Control in Dictatorships.* Technical report. Discussion Paper, Center for Interdisciplinary Economics.

Barrett, Christopher B, ed. 2013. *Food Security and Sociopolitical Stability.* Oxford University Press.

Bazzi, Samuel, and Christopher Blattman. 2014. "Economic Shocks and Conflict: Evidence from Commodity Prices." *American Economic Journal: Macroeconomics* 6 (4): 1–38.

Belkin, Aaron, and Evan Schofer. 2003. "Toward a Structural Understanding of Coup Risk." *Journal of Conflict Resolution* 47 (5): 594–620.

Besley, Timothy, and Masayuki Kudamatsu. 2008. "Making Autocracy Work." In *Institutions and Economic Performance,* edited by Elhanan Helpman. Cambridge Mass.: Harvard University Press.

Blattman, Christopher, and Edward Miguel. 2010. "Civil War." *Journal of Economic Literature* 48 (1): 3–57.

Brueckner, Markus, and Antonio Ciccone. 2011. “Rain and the Democratic Window of Opportunity.” *Econometrica* 79 (3): 923–947.

Brueckner, Markus, Antonio Ciccone, and Andrea Tesei. 2012. “Oil Price Shocks, Income, and Democracy.” *Review of Economics and Statistics* 94 (2): 389–399.

Bueno de Mesquita, Bruce, Alastair Smith, Randolph M. Siverson, and James D. Morrow. 2005. *The Logic of Political Survival.* The MIT Press.

Burke, Marshall, Solomon M. Hsiang, and Edward Miguel. 2014. *Climate and Conflict.* NBER Working Paper 20598.

Burke, Paul J, and Andrew Leigh. 2010. “Do Output Contractions Trigger Democratic Change?” *American Economic Journal: Macroeconomics* 2 (4): 124–157.

Casper, Brett A., and Scott A. Tyson. 2014. “Popular Protest and Elite Coordination in a Coup d’état.” *The Journal of Politics* 76 (02): 548–564.

Cheibub, Jose Antonio, Jennifer Gandhi, and James Raymond Vreeland. 2010. “Democracy and Dictatorship Revisited.” *Public Choice* 143 (1-2): 67–101.

Ciccone, Antonio. 2011. “Economic Shocks and Civil Conflict: A Comment.” *American Economic Journal: Applied Economics* 3 (4): 215–227.

Collier, Paul, and Benedikt Goderis. 2012. “Commodity Prices and Growth: An Empirical Investigation.” *European Economic Review* 56:1241–1260.

Dell, Melissa, Benjamin F. Jones, and Benjamin A. Olken. 2014. “What Do We Learn from the Wheather? The New Climate-Economy Literature.” *Journal of Economic Literature* 52 (3): 740–798.

Galetovic, Alexander, and Ricardo Sanhueza. 2000. “Citizens, Autocrats, and Plotters: A Model and New Evidence on Coups D’Etat.” *Economics & Politics* 12 (2): 183–204.

Gassebner, Martin, Jerg Gutmann, and Stefan Voigt. 2016. “When to Expect a Coup d’État? An Extreme Bounds Analysis of Coup Determinants.” *Public Choice* 169:293–313.

Geddes, Barbara. 1999. “What Do We Know About Democratization After Twenty Years?” *Annual Review of Political Science* 2 (1): 115–144.

Gilli, Mario, and Yuan Li. 2015. "Coups, Revolutions and Efficient Policies in Autocracies." *European Journal of Political Economy* 39:109–124.

Gleditsch, Nils P., Peter Wallensteen, Mikael Eriksson, Margareta Sollenberg, and Håvard Strand. 2002. "Armed Conflict 1946-2001: A New Dataset." *Journal of Peace Research* 39 (5): 615–637.

Harari, Mariaflavia, and Eliana La Ferrara. 2014. *Conflict, Climate and Cells: A Disaggregated Analysis.* CEPR Discussion Paper No. DP9277.

Hendrix, Cullen S., and Stephan Haggard. 2015. "Global Food Prices, Regime Type, and Urban Unrest in the Developing World." *Journal of Peace Research* 52 (2): 143–157.

Hendrix, Cullen S., and Idean Salehyan. 2012. "Climate Change, Rainfall, and Social Conflict in Africa." *Journal of Peace Research* 49 (1): 35–50.

Johnson, Jaclyn, and Clayton L. Thyne. 2016. "Squeaky Wheels and Troop Loyalty: How Domestic Prostests Influence Coups d'états." *Journal of Conflict Resolution* Doi: 10.1177/0022002716654742.

Kim, Nam Kyu. 2014. "Revisiting Economic Shocks and Coups." First published on February 18, 2014, *Journal of Conflict Resolution:* 1–29.

Kuran, Timur. 1989. "Sparks and Prairie Fires: A Theory of Unanticipated Political Revolution." *Public Choice* 61 (1): 41–74.

Londregan, John B., and Keith T. Poole. 1990. "Poverty, the Coup Trap, and the Seizure of Executive Power." *World Politics* 42 (2): 151–183.

Marinov, Nikolay, and Hein Goemans. 2014. "Coups and Democracy." *British Journal of Political Science* 44 (04): 799–825.

Marshall, Monty G., Ted R. Gurr, and Keith Jaggers. 2014. *Polity IV Project: Political Regime Characteristics and Transitions, 1800–2013: Dataset Users' Manual.* Vienna: Center for Systemic Peace.

Marshall, Monty G., and Donna Ramsey Marshall. 2015. *Coup D'État Events, 1946 - 2014: Codeboook.* Center for Systemic Peace.

Mehlum, Halvor, Karl Moene, and Ragnar Torvik. 2006. "Institutions and the Resource Curse." *The Economic Journal* 116:1–20.

Miguel, Edward, and Shanker Satyanath. 2011. "Re-Examining Economic Shocks and Civil Conflict." *American Economic Journal: Applied Economics* 3 (4): 228–232.

Miguel, Edward, Shanker Satyanath, and Ernest Sergenti. 2004. "Economic Shocks and Civil Conflict: An Instrumental Variables Approach." *Journal of Political Economy* 112 (4): 725–753.

Miller, Michael K. 2012. "Economic Development, Violent Leader Removal, and Democratization." *American Journal of Political Science* 56 (4): 1002–1020.

Powell, Jonathan. 2012. "Determinants of the Attempting and Outcome of Coups D'État." *Journal of Conflict Resolution* 56 (6): 1017–1040.

Powell, Jonathan M, and Clayton L Thyne. 2011. "Global Instances of Coups from 1950 to 2010 A new dataset." *Journal of Peace Research* 48 (2): 249–259.

Raleigh, Clionadh, Hyun Jin Choi, and Dominic Kniveton. 2015. "The Devil is in the Details: An Investigation of the Relationships between Conflict, Food Price and Climate across Africa." *Global Environmental Change* 32:187–199.

Raleigh, Clionadh, Andrew Linke, and Hegre Havard. 2010. "Introducing ACLED: An Armed Conflict Location and Event Dataset." *Journal of Peace Research* 47 (5): 651–660.

Salehyan, Idean, and Cullen S Hendrix. 2014. *Social Conflict in Analysis Database Version 3.1: Codebook and Coding Procedures.*

Salehyan, Idean, Cullen S Hendrix, Jesse Hamner, Christina Case, Christopher Linebarger, Emily Stull, and Jennifer Williams. 2012. "Social Conflict in Africa: A New Database." *International Interactions* 38 (4): 503–511.

Smith, Todd G. 2014. "Feeding Unrest: Disentangling the Causal Relationship between Food Price Shocks and Sociopolitical Conflict in Urban Africa." *Journal of Peace Research* 51 (6): 679–695.

Stock, James H., and Motohiro Yogo. 2002. *Testing for Weak Instruments in Linear Regressions.* NBER Technical Working Paper 284.

Svolik, Milan W. 2013. "Contracting on Violence The Moral Hazard in Authoritarian Repression and Military Intervention in Politics." *Journal of Conflict Resolution* 57 (5): 765–794.

Thyne, Clayton L. 2010. "Supporter of Stability or Agent of Agitation? The Effect of US Foreign Policy on Coups in Latin America, 1960–99." *Journal of Peace Research* 47 (4): 449–461.

Thyne, Clayton L., and Jonathan M. Powell. 2014. "Coup d'État or Coup d'Autocracy? How Coups Impact Democratization, 1950–2008." *Foreign Policy Analysis* Doi: 10.1111/fpa.12046.

Tullock, Gordon. 1971. "The Paradox of Revolution." *Public Choice* 11 (1): 89–99.

van Weezel, Stijn. 2016. "Food Imports, International Prices, and Violence in Africa." *Oxford Economic Papers* Doi: 10.1093/oep/gpw015.

Wig, Tore, and Espen G. Rød. 2014. "Cues to Coup Plotters: Elections as Coup Triggers in Dictatorships." *Journal of Conflict Resolution* 60 (5): 787–812.

World Bank. 2015. *World Development Indicators.* Washington, DC: The World Bank Group.

Appendix

Table A.1: Summary Statistics

	Obs	Mean	Std. Dev.	Min	Max
Baseline Model					
Coup	682	0.056	0.23	0	1
Urban protest, logs	682	0.464	1.063	0	6.004
GDP p.c. growth	682	0.007	0.059	-0.502	0.371
Drought	682	0.17	0.376	0	1
Drought × Western Africa dummy	682	0.081	0.272	0	1
Drought × Eastern Africa dummy	682	0.004	0.066	0	1
Drought × Middle Africa dummy	682	0.037	0.188	0	1
Drought × Southern Africa dummy	682	0.048	0.215	0	1
Extractive price shock	682	-0.008	0.891	-6.208	7.529
GDP p.c., logs	682	6.335	0.9424	4.733	9.022
Polity2	682	-0.0762	5.419	-10	9
Military expenditure, logs	585	22.384	3.0127	4.0395	30.912
Alternative Coup & Protest Measures					
Coupcount	682	0.069	0.315	0	4
Coup, CSP	682	0.073	0.261	0	1
Successful coup, CSP	682	0.023	0.151	0	1
Failed coup, CSP	682	0.05	0.218	0	1
Successful coup, PT	682	0.026	0.16	0	1
Failed coup, PT	682	0.029	0.169	0	1
Rural protest, logs	682	2.335	1.728	0	7.249
Protest, not weighted (logs)	682	2.424	1.74	0	6.582
Protest, near capital (logs)	682	0.95	0.97	0	4.82
Alternative Instruments					
Drought × agricultural employment	538	8.285	20.661	0	83
Regime Type					
Initial Polity2 value	682	-5.389	4.442	-10	8
Initial military regime	682	0.554	0.497	0	1
Initial civil regime	682	0.419	0.494	0	1
Initial royal regime	682	0.026	0.160	0	1

Table A.2: List of Sample Countries

Western Africa	Eastern Africa	Middle Africa	Southern Africa
Benin	Burundi	Angola	Botswana
Burkina Faso	Ethiopia	Cameroon	Lesotho
Cote d'Ivoire	Kenya	Central African Republic	Namibia
The Gambia	Madagascar	Chad	South Africa
Ghana	Malawi	DR Congo	Swaziland
Guinea	Mozambique	Republic Congo	
Guinea-Bissau	Rwanda	Gabon	
Liberia	Tanzania	Sudan	
Mali	Uganda		
Mauritania	Zambia		
Niger	Zimbabwe		
Nigeria			
Senegal			
Sierra Leone			
Togo			

Table A.3: 2SLS, Alternative Protest Variables and Instruments, First-Stage Results for *Rural Protest* and *GDP p.c. Growth*

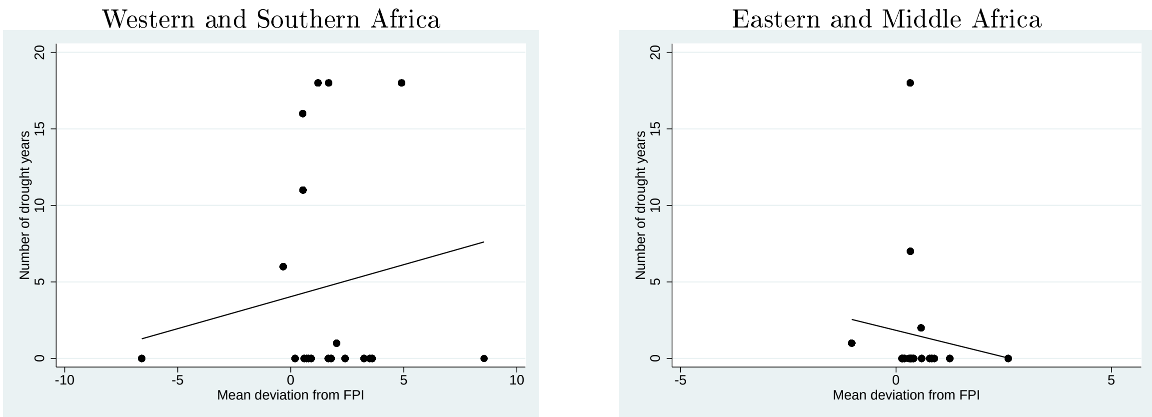
Dep. car. in the structural regression: Coup$_t$	Model (1)	Model (2)	Model (3)	Model (4)	Model (5)
	C. First stage for *rural protest*$_t$				
Drought$_t$	1.167***				
	(0.342)				
Drought$_t\times$ Eastern Africa	-1.512**				
	(0.687)				
Drought$_t\times$ Middle Africa	-2.0999***				
	(0.635)				
Drought$_t\times$ Southern Africa	-0.917*				
	(0.507)				
Extractive price shock$_t$	-0.0947*				
	(0.0518)				
Country & year FE	Y				
F-statistic on excl. IVs	3.53				
	D. First stage for *GDP p.c. growth*$_t$				
Drought$_t$	-0.00203	-0.00203	-0.00203		
	(0.00757)	(0.00757)	(0.00757)		
Drought$_t\times$ Eastern Africa	-0.0236	-0.0236	-0.0236		
	(0.0267)	(0.0267)	(0.0267)		
Drought$_t\times$ Middle Africa	-0.00789	-0.00789	-0.00789		
	(0.0206)	(0.0206)	(0.0206)		
Drought$_t\times$ Southern Africa	0.00859	0.00859	0.00859		
	(0.0116)	(0.0116)	(0.0116)		
Drought$_t\times$ agricultural employment				-0.000255	0.0000199
				0.000334	(0.000197)
Drought$_t\times$ agri. emp.$\times$ Eastern Africa					-0.000466
					(0.000470)
Drought$_t\times$ agri. emp.$\times$ Southern Africa					0.000274
					(0.000627)
Extractive price shock$_t$	-0.00507**	-0.00507**	-0.00507**	-0.00492*	-0.00486*
	(0.00230)	(0.00230)	(0.00230)	(0.00253)	(0.00254)
Country & year FE	Y	Y	Y	Y	Y
F-statistic on excl. IVs	1.47	1.47	1.47	2.20	1.32

Note: First-stage results from table 3 for *rural protest, logs* and *GDP p.c. Growth*. *Rural protest, logs* is the logged duration of protest events located in areas with less than 100 individuals per km^2. *Agricultural employment* is the share of agricultural employment in total employment in 1990. Significance levels: *** p<0.01, ** p<0.05, * p<0.1

Table A.4: Alternative Measures of Coups and Alternative Sample Criteria, First-Stage Results for *urban protest* and *GDP p.c. growth*

	First-stage results for model (1) to (5)		First-stage results for model (6)	
Dependent variable in the first-stage regression	Urban protest$_t$	GDP p.c. growth$_t$	Urban protest$_t$	GDP p.c. growth$_t$
	(1)	(2)	(3)	(4)
Drought$_t$	2.651***	-0.00203	2.920***	0.00622
	(0.246)	(0.00757)	(0.142)	(0.00764)
Drought$_t\times$ Eastern Africa	-2.299***	-0.0236	-2.563***	-0.0268
	(0.579)	(0.0267)	(0.549)	(0.0249)
Drought$_t\times$ Middle Africa	-2.610***	-0.00789		
	(0.260)	(0.0206)		
Drought$_t\times$ Southern Africa	-2.404***	0.00859	-2.694***	-0.00216
	(0.325)	(0.0116)	(0.258)	(0.0115)
Extractive price shock$_t$	-0.0210	-0.00507**	-0.0140	-0.00207
	(0.0250)	(0.00230)	(0.0304)	0.(00237)
Country & year FE	Y	Y	Y	Y
F-statistic on excl. IVs	27.64	1.47	106.07	0.6685

Note: First-stage results from table 4 for *urban protest, logs* and *GDP p.c. growth*. Columns (1) and (2) report the first-stage results for model (1) to (5) in table 4, columns (3) and (4) report the first-stage results for model (6) in table 5. Significance levels: *** p<0.01, ** p<0.05, * p<0.1



Note: FPI: Food price index from van Weezel (2016), country-specific average over monthly standardized deviation from mean FPI. *Number of drought years:* Total number of drought years experienced by a country during the sample period (1990-2007). Unit of observation is the country.

Figure A.1: Drought Years and Food Prices across African Regions