UNIVERSITÀ CATTOLICA del Sacro Cuore

DIPARTIMENTO DI POLITICA ECONOMICA

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Quaderno n. 57/January 2026

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ISBN digital edition (PDF): 978-88-343-6218-1

www.vitaepensiero.it

PROGETTO DI RICERCA DI RILEVANTE INTERESSE NAZIONALE 2022

Climate Change, Violent Conflicts and Welfare: A Multi-Scale Investigation of Causal Pathways in Different Institutional Contexts

PNRR per la Missione 4, Componente 2, investimento 1.1. Avviso 104/2022 Finanziato dall'Unione europea – Next Generation EU. CClimate-Conflicts – Prot. 2022RSZW83 – CUP J53D23005870008

Abstract

We analyse to what extent land-related institutional settings affect the likelihood of communal violence in Sub-Saharan Africa and whether this relationship is conditioned by climate variability. Using a country–year panel covering the period 1990–2024, we focus on the occurrence of communal violence and examine the role of legal transparency and predictable enforcement of laws. The empirical analysis relies on a panel probit model for binary outcomes, controlling for socio-economic characteristics, land-use patterns, demographic pressure, and conflict persistence. The results show that higher levels of legal transparency and more predictable enforcement are consistently associated with a significantly lower likelihood of communal violence. This relationship proves robust across alternative specifications and sample restrictions. To address potential endogeneity in institutional quality, we implement a set of complementary strategies to account for unobserved heterogeneity, while exploiting early post-independence institutional conditions to mitigate concerns related to reverse causality. These checks support the robustness of the baseline association. Climate variability does not emerge as an independent driver of communal violence. Instead, drought acts as a threat multiplier by conditionally weakening the conflict-mitigating effect of legal institutions. Interaction effects indicate that while improvements in institutional quality substantially reduce the probability of communal violence under normal climatic conditions, this stabilizing effect progressively diminishes as drought severity increases and becomes negligible under severe drought. Therefore, as drought severity increases, the mitigating role of institutions progressively weakens. Overall, the findings highlight the central role of legal transparency and predictable enforcement in managing land-related tensions, while showing that their effectiveness is contingent on environmental stress.

JEL classification: D74, O13, Q54
Keywords: communal violence, land institutional settings, climate shock, conflicts, Africa.

Acknowledgments
We gratefully thank the participants of the XXIII Jan Tinbergen European Peace Science Conference of the Network of European Peace Scientists (NEPS), the 65nd Annual Conference of the Italian Economic Association (SIE), the IX International Conference of the Italian Association of Development Economists, and the workshop Conflicts, Violence and the Economy (Università Cattolica del Sacro Cuore, 2024) for useful comments on earlier versions of this article. The authors acknowledge support by the Italian Ministero dell'Istruzione, dell'Università e della Ricerca, PRIN-2022 project 2022RSZW83 "Climate change, violent conflicts and welfare: A multi-scale investigation of causal pathways in different institutional contexts (CCLIMATE-CONFLICTS)" (principal investigator: Raul Caruso). Data supporting the findings of this study are available from the corresponding author upon request.

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1 Introduction

Violent confrontations between local communities over land and natural resources remain a salient feature of non-state conflict dynamics in Sub-Saharan Africa (Davies, Pettersson, Sollenberg & Öberg, 2025). Communal violence refers to a form of non-state conflict involving the use of armed force between organized groups that are not state or governmental actors, where the opposing parties are defined primarily by communal or identity-based characteristics such as ethnicity, religion, tribe, or similar markers. Such violence entails lethal confrontations between these groups, even when the actors are not formally structured as militias or armed organizations. In many contexts, communal violence emerges from competition over land and other vital resources, particularly where livelihoods depend heavily on agriculture and land-based production (Binswanger, Deininger & Feder, 2017). A growing literature in conflict studies documents how land-related grievances can both trigger violence and, when effectively governed, mitigate it (Madrigal, Cuesta & Somerville, 2025). Yet, despite increasing attention to environmental and climatic stressors, less is known about how institutional arrangements governing land shape the relationship between resource competition and communal violence.

Disputes over grazing areas, farmland boundaries, and access to common-pool resources are more likely to escalate when legal systems fail to provide clear, transparent, and enforceable rules (Deininger, Selod & Burns, 2012). Weak or unpredictable enforcement increases uncertainty over land claims, lowers the costs of opportunistic behaviour, and undermines non-violent dispute resolution (North, 1990; Williamson, 2000). By contrast, legal transparency and predictable enforcement reduce information asymmetries and constrain discretionary power, shaping incentives for cooperation among land-dependent groups (Acemoglu, Johnson & Robinson, 2001; Glaeser, La Porta, Lopez-de-Silanes & Shleifer, 2004). These mechanisms underscore the central role of institutional quality in structuring the risks of communal conflict.

These institutional dynamics are particularly pronounced in Sub-Saharan Africa, where land constitutes a primary productive asset and where statutory and customary tenure systems often coexist. In such settings, institutions governing land access and use are not merely administrative arrangements but key determinants of group relations and conflict dynamics (Eck, 2014). When legal frameworks are weak or contested, competition over land can readily become politicized along ethnic, communal, or local identity lines, increasing the likelihood of violent mobilization.

At the same time, many land-dependent societies are increasingly exposed to climate variability. For instance, herder–farmer disputes in Nigeria, water tensions in South Sudan, and grazing conflicts in Kenya often intensify under climate stress, illustrating how resource scarcity can exacerbate local disputes (McGuirk & Nunn, 2025). Droughts and other climate shocks can disrupt agricultural production, alter mobility patterns, and heighten competition over land and water (Fjelde & Von Uexkull, 2012; Raleigh & Kniveton, 2012). Importantly, climate variability does not mechanically translate into violence. Consistent with a large literature on conflict, economic and political conditions shape whether environmental stress results in violent outcomes (Buhaug, 2015; Koubi, 2019; Miguel, Satyanath & Sergenti, 2004). In this perspective, climate-related shocks operate as conditional stressors whose conflict potential depends on the institutional environment in which they occur (McGuirk & Nunn, 2020).

This paper advances the literature by examining whether and how land-related institutional quality influences the relationship between climate variability and communal violence. We focus on legal transparency and predictability as core institutional features shaping land governance, and argue that these characteristics influence how climate-induced economic shocks translate into communal conflict. Rather than treating climate variability as an independent driver of violence, we conceptualize it as a threat multiplier that affects the role of institutions under conditions of environmental stress.

Empirically, we analyse a country–year panel covering Sub-Saharan Africa over the period 1990–2024. We estimate panel probit models in which the occurrence of communal violence is explained by institutional quality, land-use patterns, socio-economic conditions, and conflict persistence, and explicitly interact institutional quality with drought indicators. This design allows us to assess whether the effect of institutions on communal violence varies systematically with climate stress.

Our findings show that higher levels of legal transparency and predictable enforcement are associated with a lower likelihood of communal violence. In fact, climate variability alone does not increase the probability of conflict. Instead, our results show that environmental stress affects the risk of communal violence through its interaction with the conflict-mitigating role of land-related institutional quality. In particular, as drought severity increases, the mitigating role of institutions becomes ineffective.

The remainder of the paper is organized as follows. Section 2 reviews the relevant literature, while Section 3 outlines the empirical strategy. Section 4 describes the data. Section 5 presents the main results and discusses a set of robustness checks, including additional analyses related to potential endogeneity concerns. Section 6 concludes.

2 *Land, Institutions, and Communal Violence: Related Literature*

A large body of economic literature identifies the rule of law—defined by the transparency, consistency, and predictability of legal rules and their enforcement—as a key determinant of natural resource governance, including land. Predictable enforcement lowers transaction costs, constrains opportunistic behaviour, and supports long-term investment by securing property and use rights (North, 1990; Williamson, 2000). These mechanisms are particularly salient for land, where weak or opaque legal systems increase uncertainty over land claims, encourage informal dispute resolution, and often result in inefficient land use or conflict. In contrast, transparent rules and credible enforcement create a stable institutional environment in which land can be allocated, transferred, and managed more efficiently (Acemoglu et al., 2001; Glaeser et al., 2004).

Security of land tenure depends not only on the formal definition of property rights but critically on the credibility and predictability of their enforcement. Secure tenure is robustly associated with higher investment, greater land productivity, and improved land management (Deininger & Feder, 2009; Goldstein & Udry, 2008). Micro- and cross-country evidence shows that stronger property-rights institutions—often proxied by rule of law or contract enforcement indicators—are correlated with more efficient land use and lower land degradation (Ali, Deininger & Goldstein, 2014; Field, 2007). This literature emphasizes that formal land titling alone is insufficient when enforcement is weak or discretionary: without credible legal institutions, titles may fail to provide de facto security, undermining incentives for long-term investment and sustainable land management.

Legal institutions also shape how competing claims over land and resources are resolved. Transparency in legal and administrative processes reduces information asymmetries and limits discretionary power in land allocation and dispute resolution (Deininger et al., 2012).

These institutional mechanisms are especially consequential in agrarian economies, where land constitutes the primary productive asset and livelihoods are directly exposed to environmental variability. In much of Sub-Saharan Africa, agriculture remains predominantly rain-fed and land-intensive, and land relations are shaped by a mix of statutory and customary norms, with profound implications for access, exclusion, and investment (Binswanger et al., 2017). In such contexts, weak or unpredictable legal enforcement does not merely affect investment incentives but can directly shape patterns of resource competition and conflict, particularly under climate-induced stress. Where formal enforcement is absent or unreliable, land governance often relies on informal institutions or ad hoc dispute resolution mechanisms. While these arrangements can partially substitute

for formal systems, they are typically limited in scale and vulnerable to elite capture, especially when resource scarcity intensifies. In these settings, groups whose livelihoods are directly tied to land and renewable resources are particularly exposed to institutional failures in resource allocation and dispute resolution (Blattman, Hartman & Blair, 2014).

We focus specifically on communal violence, as its nature makes it particularly sensitive to changes in access to land and other vital resources. Communal groups are not permanently organized for combat but mobilize based on shared identity, such as ethnic or clan affiliation, when they perceive threats to their identity or to the livelihoods they depend on. Collective identity, social cohesion, and group affiliation can act as powerful mechanisms that amplify tensions and trigger outbreaks of violence (Esteban, Mayoral & Ray, 2012; Sambanis & Shayo, 2013), particularly when groups face resource scarcity (Homer-Dixon, 2010) or social marginalization (Hillesund, 2019). Competition over grazing land, water, and other renewable resources can rapidly escalate into communal violence, with looting, raids, and inter-community clashes serving as well-documented examples (Detges, 2014; Döring, 2020; Scheffran, Ide & Schilling, 2017).

Violent communal conflicts between identity-based groups pose a significant threat to human security and development, causing large-scale fatalities (Krause, 2020), severe disruptions to livelihoods, and further destabilization of conflict-affected societies (Brosché, 2023). Although they often emerge in countries already affected by civil war, they are not an inevitable consequence of broader armed conflict (van Baalen, 2024). While these conflicts frequently originate at the local level, they can escalate and spill over into surrounding areas (Balestri & Maggioni, 2017), generating cycles of retaliation and contributing to wider patterns of instability (Van Weezel, 2019).

Our analysis aligns with a growing literature emphasizing the role of institutions in shaping the relationship between environmental shocks, resource access, and communal conflict in Africa. Eck, 2014 shows that communal land conflicts are more likely where institutional uncertainty arises from competing legal authorities, while Wig and Kromrey, 2018 finds that well-developed customary institutions can reduce conflict by facilitating credible non-violent bargaining at the group level. Complementing these perspectives, Petrova, 2022 demonstrates that the effect of flood disasters on communal violence depends on trust in local political and judicial institutions, suggesting that environmental shocks translate into conflict only under specific institutional conditions.

Building on this literature, we argue that institutional quality shapes both how resources are allocated and how disputes are managed. By concentrating on communal violence, we can more precisely isolate the effects of weak or unpredictable legal systems on the management of land and resource-based livelihoods, while accounting for the amplifying role of climate variability and environmental stress. At the same time, climate variability can heighten livelihood insecurity, mobility pressures, and competition over resources—particularly among vulnerable, land-dependent groups—thereby increasing the risk of communal violence (Balestri & Caruso, 2024; Döring, 2020; Fjelde & Von Uexkull, 2012; Raleigh & Kniveton, 2012; Van Weezel, 2019). Consistent with a large and growing literature on conflict, we follow (Miguel et al., 2004) in emphasizing that economic shocks—often triggered by environmental stressors—may drive conflict[1] In our analysis, we emphasize that institutional factors can shape how climate-induced economic shocks translate into conflict, potentially amplifying or mitigating their effects. By employing a time-varying, cross-country measure of legal transparency and predictability, our approach allows us to examine how climate-related stressors and institutional quality jointly influence the risk of communal violence across Sub-Saharan Africa over a broad period.

[1] See Romano et al., 2025 for a recent review of the climate–agrifood–conflict nexus and its multiple pathways.

3 Research question and methods

This paper investigates whether and how land-related institutional settings affect the likelihood of communal violence in Sub-Saharan Africa, and whether this relationship depends on climatic conditions. Specifically, we ask whether greater legal transparency and predictable enforcement of laws reduce the risk of communal violence, and whether their conflict-mitigating effect is weakened or reinforced by climate variability.

We argue that in settings characterized by widespread reliance on land-based livelihoods and frequent environmental stress, institutions play a central but potentially conditional role. While strong institutional arrangements may reduce uncertainty, grievances, and opportunistic violence, climate shocks—particularly droughts—may intensify competition over scarce resources and erode the effectiveness of institutional conflict-management mechanisms.

To address these issues, we combine country-year data on communal violence, institutional quality, land use, and climate variability for Sub-Saharan Africa over the period 1990–2024. We estimate a panel probit model in which the dependent variable captures the occurrence of communal violence at the country-year level.

The baseline specification is given by:

$$P(confl_{i,t} = 1 \mid X_{i,t}, u_i) = \Phi\Big(\alpha + \beta_1 InstSettings_{i,t-1} + \beta_2 SocioEcon_{i,t-1} + \beta_3 LandUse_{i,t-1} + \beta_4 PastConfl_{i,t-1} + u_i\Big) \tag{1}$$

where the dependent variable indicates whether at least one episode of communal violence occurred in country i at time t. *InstSettings* captures land-related institutional settings, *SocioEcon* includes economic and demographic controls, *LandUse* measures specific land uses—namely, the share of pastureland and forests—and *PastConfl* accounts for the lagged occurrence of communal violence. All covariates are lagged by one year to mitigate concerns related to reverse causality.

Additional considerations concern the structure of the data and the choice of the panel estimator. Our dataset consists of an almost perfectly balanced panel of 43 countries over 32 years[2], with a binary dependent variable indicating the occurrence of a relatively rare event[3]. These features led us to opt for a random effects probit model rather than a fixed effects specification. Several methodological considerations motivated this choice. First, fixed effects in limited dependent variable models are known to suffer from the incidental parameter problem (IPP), particularly when the number of cross-sectional units (N) is large relative to the number of time periods (T). Although our panel is relatively long (T = 32), IPP remains a concern given the modest number of countries (N = 43) and the binary nature of the outcome. As Greene (2004) notes, incidental parameter bias can persist in non-linear panel models even for moderate T, potentially distorting estimates of structural parameters. Second, a number of countries in our sample never experience the event of interest (i.e., the dependent variable is always zero). In fixed effects probit models, such units are automatically dropped due to the lack of within-unit variation, which can lead to sample selection bias and reduce the representativeness of the results. Third, random effects models allow us to retain all units and leverage the full sample, provided that the unobserved unit-specific effects are approximately orthogonal to the covariates. To assess the plausibility of this assumption, we conducted a Hausman-type specification test adapted for non-linear panel models (Wooldridge, 2019), using country-specific time averages of the covariates. Overall, the results provide reasonable support for the orthogonality assumption, suggesting that a random effects specification is a suitable choice for our panel.

Finally, because observations within the same country are likely correlated over time, we clustered standard errors at the country level to correct for intra-unit correlation and avoid downward-biased standard errors.

[2] Two countries are observed for a shorter period, namely Eritrea (since 1993) and South Sudan (since 2011)

[3] Events of communal violence are recorded in 294 country/year observations out of 1,481 total.

This combination of random effects estimation and clustered standard errors allows us to mitigate incidental parameter bias, avoid unnecessary loss of information, and obtain robust inference.

In a second stage of the analysis, we examine the role of climate variability on the relationship between institutional quality and communal violence. To this end, we explicitly interact our measure of land-related institutional settings with indicators of drought conditions. This interaction allows us to identify whether the conflict-mitigating effect of stronger institutions is conditional on climatic conditions and to quantify how adverse or favourable environmental shocks may amplify or attenuate the effectiveness of institutional mechanisms in preventing communal violence.

Building on the baseline models, we extend the specification as follows:

$$P(d\_communal_{i,t} = 1 \mid X_{i,t}, u_i) = \Phi\Big(\alpha + \beta_1 InstSettings_{i,t-1} + \beta_2 (InstSettings_{i,t-1} \times ClimVar_{i,t-1}) + \beta_3 SocioEcon_{i,t-1} + \beta_4 LandUse_{i,t-1} + \beta_5 PastConfl_{i,t-1} + u_i\Big) \tag{2}$$

A final step of the analysis explores potential endogeneity concerns in the relationship between institutional quality and communal violence. Although the type of violence we consider—namely, communal violence—typically involves a low level of organisation and affects only a few social groups, it often occurs in specific areas of the country and is therefore unlikely to have a substantial shaping effect on national legal institutions. Nevertheless, the level of transparency and predictable enforceability of laws could be endogenous to the level of violence experienced by a given country. In addition, endogeneity concerns may also arise from omitted variables and other unobserved factors, potentially biasing our estimates.

To mitigate potential sources of endogeneity, we adopt a twofold strategy. First, we apply a Mundlak-style correlated random effects specification to account for unobserved time-invariant heterogeneity (Mundlak, 1978). This method mitigates potential bias due to omitted, time-invariant country-specific factors, improving the reliability of the estimated relationships. Second, we implement a Control Function approach to assess the robustness of our results to potential reverse causality. In this analysis, contemporary land-related institutional settings are instrumented using a deep lag reflecting early post-independence conditions in each country. The choice of this instrument is motivated by the idea that institutional arrangements established in the immediate post-independence period capture deep features of state formation that are persistent over time and plausibly related to contemporary levels of legal transparency and enforceability. Consistent with the historical evidence on the persistence of foundational institutions (Acemoglu et al., 2001; Besley & Persson, 2011), these early institutional conditions are expected to be correlated with contemporary land-related institutional settings, but due to the substantial temporal distance between this formative period—typically in the 1950s or 1960s—and the episodes of communal violence in our sample (1990–2024), any direct effect on contemporary violence cannot be established.

The Control Function approach is implemented through the following system of equations.

$$\begin{aligned} InstSettings_{i,t-1} &= \pi_0 + \pi_1 EarlyInst_i + \pi_2 SocioEcon_{i,t-1} \\ &\quad + \pi_3 LandUse_{i,t-1} + \pi_4 PastConfl_{i,t-1} + \nu_{i,t}, \\ P(confl_{i,t} = 1 \mid X_{i,t}, u_i) &= \Phi\Big(\alpha + \beta_1 InstSettings_{i,t-1} \\ &\quad + \beta_2 (InstSettings_{i,t-1} \times ClimVar_{i,t-1}) + \gamma residuals_{i,t} \\ &\quad + \beta_3 SocioEcon_{i,t-1} + \beta_4 LandUse_{i,t-1} + \beta_5 PastConfl_{i,t-1} + u_i\Big) \end{aligned} \tag{3}$$

where $EarlyInst_i$ denotes the average level of land-related institutional settings during the first five years of

independence, $residuals_{i,t}$ are the first-stage residuals capturing possible feedback effects, and u_i represents unobserved, time-invariant country-specific effects.

4 Data description

4.1 Communal violence

Communal conflicts are violent events between non-state groups organized around shared communal identities - such as ethnicity, religion, or tribe. These conflicts often occur when groups perceive threats to their identity, resources, or status, leading to escalations that can even result in significant loss of life and displacement. We relied on data collected from the UCDP Georeferenced Event Dataset (UCDP-GED) (Sundberg & Melander, 2013) and the Non-State Conflict Dataset v.24.1 (Sundberg, Eck & Kreutz, 2012) to identify their occurrence. For each non-state conflict[4], information is provided on the actors involved and their level of organization. We have therefore selected those conflicts where the actors-dyad is composed of informal groups that are not permanently organized for combat, but who at times organize themselves along communal identity lines to engage in fighting. Subsequently, we selected all the violent events associated with these dyads. The outbreak of individual events of communal violence is operationalized as a dichotomous variable, taking the value of 1 if an event is recorded in a given country/year, 0 otherwise.

Since 1990, communal violence has been documented in 21 countries in Sub-Saharan Africa, totaling 4,585 reported events [5]. In our sample, 19.85% of country/year observations are characterized by occurrences of communal violence. Beyond the peak of the early 1990s, driven primarily by the tragic incidents in Rwanda and Burundi, a clear upward trend in communal violence events emerges from 2008 onward. From then on, the frequency of communal violence increased sharply, becoming a significant component of overall violence in the region. Despite the overall upward trend throughout the period, there has been a slight reduction in communal violence in the last few years.

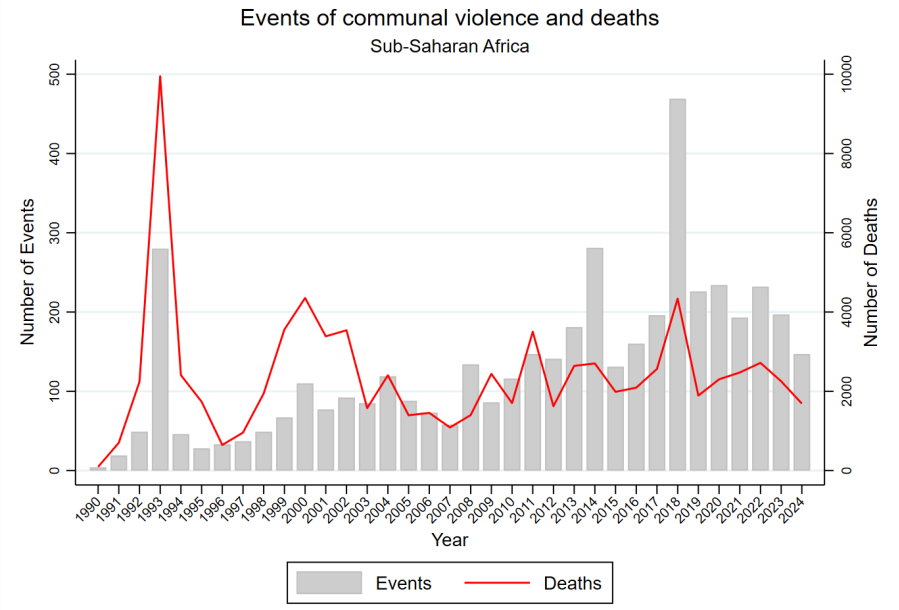

Figure 1: Occurrence of communal violence and associated deaths in Sub-Saharan Africa (1990-2024)

[4]It is defined by the Uppsala Conflict Data Program (UCDP) as "the use of armed force between two organized armed groups, neither of which is the government of a state, which results in at least 25 battle-related deaths in a year".

[5]See Appendix A for further details

Communal violence in Sub-Saharan Africa comes in various forms, depending on local contexts, resource dynamics, and historical grievances. Among the most recurring typologies, we can observe farmer–pastoralist conflicts (e.g. the violence between Fulani herders and Dogon or Mossi farming groups in Mali and Burkina Faso); ethnic or identity-based clashes (e.g. the escalating violence between ethnic Amhara, Oromo and other groups in Ethiopia); and cattle raiding and retaliatory clashes (e.g. the cyclical violence involving armed cattle rustling perpetrated by Turkana or Pokot groups in northern Kenya). Conflicts of a religious nature tend to be less frequent[6], suggesting that competition over access to resources may be a predominant factor driving communal violence in the region.

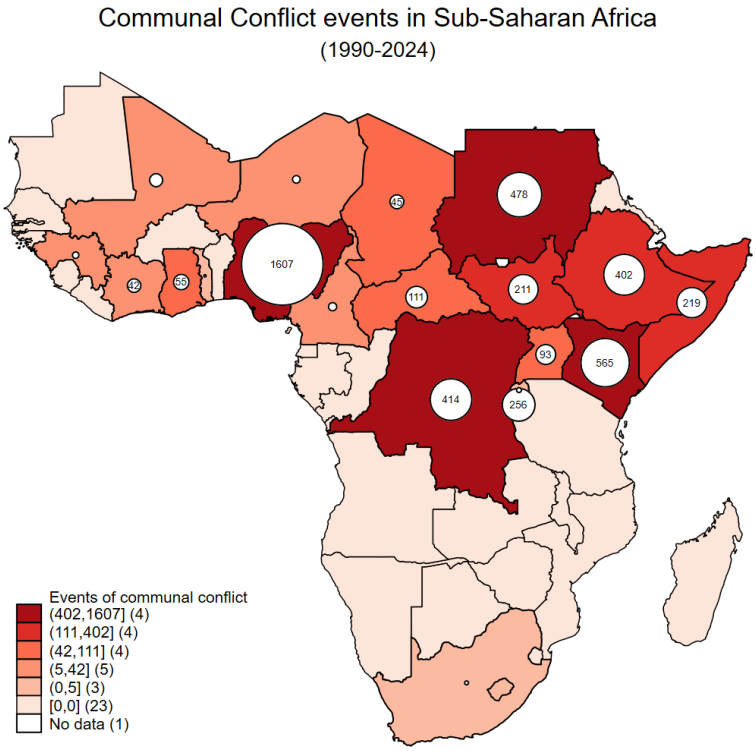

Figure 2: Total events of communal violence in Sub-Saharan Africa (1990–2024). Darker shades indicate higher incidence, and the white circles are proportional to the number of events recorded.

Figure 2 illustrates the geographic distribution of communal violence in Sub-Saharan Africa during the period of observation. The map shows that violence is concentrated in a limited set of countries, particularly along the central belt of the region.

4.2 Measuring land-related institutional settings

Empirical research on land institutions is severely constrained by the absence of standardized indicators that are comparable across countries and over time (Deininger et al., 2012). Land tenure arrangements and property rights institutions are deeply context-specific, often informal, and subject to gradual legal and de facto changes, which makes their systematic codification in cross-country longitudinal datasets particularly challenging. As a result, empirical analyses must rely on indirect measures that capture key institutional features shaping land access and governance.

This limitation is especially relevant in low-income countries, where land often represents the primary

[6]In the dataset, the only instances of communal violence with a religious dimension—specifically conflicts between Christians and Muslims—occurred in the Central African Republic and Nigeria. In the CAR, there were 111 such events recorded over 8 years, while in Nigeria, 237 events were documented across 22 years.

asset for livelihoods. In such contexts, weak legal frameworks—characterized by limited transparency and low predictability in enforcement—can exacerbate tenure insecurity, fuel disputes, and contribute to unequal land distribution. Given the lack of directly observable and comparable measures of land institutions, we therefore proxy land-related institutional quality by using an indicator that captures the transparency and predictable enforceability of a country's legal system, which constitutes a fundamental prerequisite for secure land rights and effective land governance.

This dimension is central to the analysis for several reasons. First, transparent and predictable laws are positively associated with clearer land rights, as they establish formal procedures for the acquisition, transfer, and inheritance of land (Zakout, Wehrmann & Torhonen, 2009). Legal transparency is a prerequisite for the development of effective land administration systems. Second, such frameworks promote consistent legal application, reduce discretionary decision making, improve public access to information, and limit opportunities for corruption (Deininger et al., 2012). By reducing arbitrary application and elite capture, they also constrain exclusionary practices that frequently trigger inter-group disputes (Boone, 2014). Third, transparent and predictable legal systems reduce ambiguity and perceptions of bias in land allocation or enforcement decisions. When communities perceive land decisions as opaque or politically manipulated, it heightens group grievances, often a precursor to collective violence. In contrast, predictable legal enforcement signals institutional impartiality, reducing the incentives to mobilize for violence (Fearon & Laitin, 2003).

We rely on V-Dem data v.15 (Coppedge, Gerring, Knutsen, Lindberg, Teorell, Altman et al., 2025) to gather information on impartial enforcement of the laws and, in particular, to what extent they are clear, well-publicized, coherent (consistent with each other), relatively stable from year to year, and enforced in a predictable manner across a country (Pemstein et al., 2024). The selected indicator[7] is generated through a Bayesian Item-Response Model (IRT). The outcome is a latent continuous variable that reflects the degree to which laws are transparent and predictably enforced across countries and over time (Coppedge, Gerring, Knutsen, Lindberg, Teorell, Marquardt et al., 2025). The indicator is expressed on a continuous scale ranging from negative to positive values, typically spanning approximately –3 to +3. In this framework, the zero point does not represent a substantive threshold; rather, it marks the standardized midpoint of the latent distribution derived from the Bayesian IRT model. The variable is estimated as a continuous latent trait and normalized across all country–year observations, with the global mean set to zero and a standard deviation of one.

Laws transparency and predictable enforcement exhibits substantial heterogeneity across countries, with notable temporal dynamics in a few cases (Figure 3). The average value over the period is 0.20; however, some countries experience particularly challenging situations, such as South Sudan (–1.764 on average) and Somalia (–1.603). Conversely, Namibia and Botswana occupy the opposite end of the spectrum, with comparatively high scores averaging 1.899 and 1.588, respectively. Overall, in Sub-Saharan Africa, this institutional dimension has improved over the period (1990-2024), with the regional average rising from –0.23 to 0.05, reflecting a positive trend and a substantial strengthening of law transparency and predictable enforcement.

4.3 Control variables

As outlined in the empirical estimation framework (Section 3), we include a set of control variables, drawn primarily from the existing literature, to improve the predictive accuracy of the model. Descriptive statistics for all variables, along with their data sources, are provided in Appendix A. In particular, our analysis aims to examine how the effect of institutional quality on communal outcomes may vary in response to climatic stress. To account for these potential interactions, we include a measure of climatic variability as a control.

[7] In the V-Dem dataset it is labelled "Transparent laws with predictable enforcement" (`v2cltrnslw`).

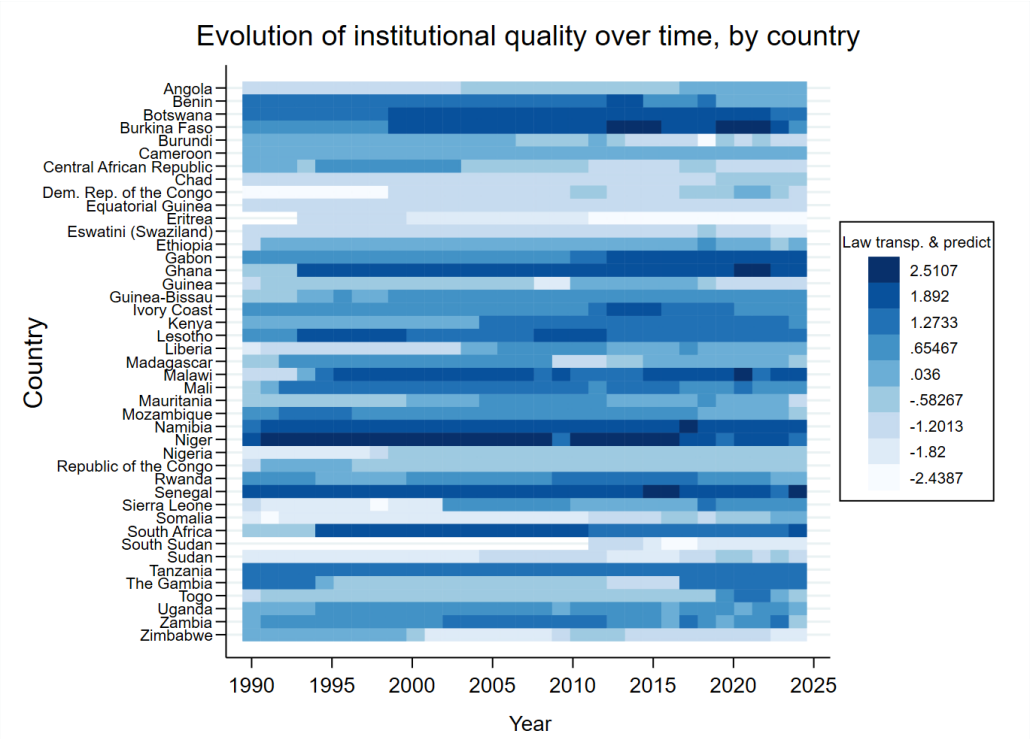

Figure 3: Evolution of institutional quality across the period 1990–2024. Darker shades indicate higher values of the transparency and predictable enforcement indicator, reflecting stronger institutional quality. Lighter tones correspond to lower values, signalling weaker institutional performance.

Climatic variability, through its interaction with institutional settings, can affect both access to land resources and their productive capacity. These changes, in turn, may alter the opportunity cost of engaging in violent actions for members of communal groups, making it a central dimension of our analysis. In particular, extreme weather events — such as prolonged droughts or heavy rainfall — can disrupt agricultural productivity and resource availability, thereby intensifying competition over land and heightening the risk of intergroup conflict. To capture these dynamics, we rely on the Standardized Precipitation Evapotranspiration Index (SPEI), which is especially well-suited for this purpose as it accounts for both precipitation and temperature, allowing for a precise measurement of water balance anomalies over multiple timescales (Beguería, Vicente-Serrano, Reig & Latorre, 2014; Vicente-Serrano et al., 2022). We apply a 12-month time-scale retrieved from the Sovereign ESG Data Portal of the World Bank. For ease of interpretation, we decompose the original SPEI index into two separate variables based on the sign of the values, using absolute values to reflect the intensity of climatic stress. Negative SPEI values, indicating drier-than-average conditions, are captured by `SPEI_dry`, while positive values, reflecting wetter-than-average conditions, are captured by `SPEI_wet`. Specifically, `SPEI_dry` equals $|\text{SPEI}|$ when $\text{SPEI} < 0$ and zero otherwise, while `SPEI_wet` equals SPEI when $\text{SPEI} > 0$ and zero otherwise. This distinction allows us to assess the potentially asymmetric effects of droughts and above-normal precipitation on communal conflict.

Over the observed period, drier conditions have been both more frequent and more severe than wetter conditions, with the average values of `SPEI_dry` and `SPEI_wet` equal to 0.544 and 0.228, respectively. This highlights the predominance of drought events across Sub-Saharan Africa and underscores the role of climate change, which is driving rising temperatures and increasingly erratic precipitation patterns (IPCC, 2022). Such drought episodes can have pronounced consequences in regions where communities rely heavily on land-based resources for their livelihoods, potentially exacerbating competition over natural resources and increasing vulnerability to socio-environmental stressors.

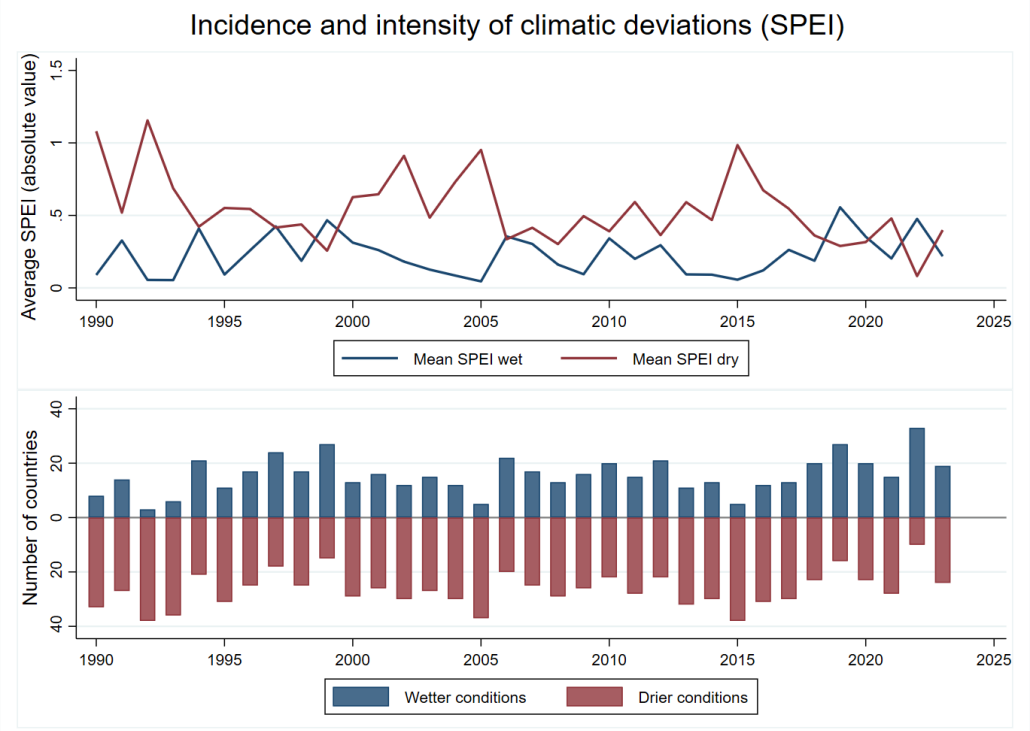

Figure 4: Incidence and intensity of climatic deviations (SPEI) across Sub-Saharan Africa. The lower panel shows the total number of countries experiencing drier-than-average conditions (negative SPEI) and wetter-than-average conditions (positive SPEI) in each year. The upper panel presents the corresponding average magnitude of SPEI deviations (absolute values) for dry and wet conditions.

When exploring the determinants of communal violence, it is crucial to consider the presence of commons—shared natural resources such as forests and pastures, as these landscapes are essential for livelihoods, identity, and access to resources. These commons can become focal points of competition, especially under conditions of scarcity, population pressure, environmental stress, or weak institutional governance (Deininger & Castagnini, 2006; Ostrom, 1990). Conflicts over access and use rights to such resources have been shown to escalate into broader communal violence, particularly where boundaries are ambiguous or customary arrangements are undermined (Raleigh & Urdal, 2007).

Pasturelands, crucial for herding communities, often become flash points of conflict with sedentary farmers, particularly where land rights are unclear or effective mediation is lacking. These clashes rank among the most common and deadly forms of communal violence in parts of the Sahel, Nigeria, Kenya, and Sudan[8]. Similarly, forests, due to their economic value and strategic importance, are frequently contested by local communities and armed groups. In the Ituri region of the Democratic Republic of the Congo, conflicts over access to timber, hunting grounds, and land for shifting cultivation have intensified recently, as armed groups and displaced populations compete with locals. Forest exploitation has become both a catalyst for violence and a source of funding for armed actors, perpetuating cycles of communal conflict.

Understanding how land use contributes to tensions helps explain the spatial and temporal dynamics of violence. To capture the presence of (potential) commons, we use the share of a country's land area covered by forests and pastures—resources that serve both as economic assets and potential sources of conflict. Given their high correlation, we include them alternately in the empirical models to avoid multicollinearity. The variables

[8]For instance, in Nigeria's Middle Belt, recurring clashes between Fulani herders and sedentary farming communities over access to pasture have resulted in thousands of deaths over the past decades. See: International Crisis Group (2017). Herders against Farmers: Nigeria's Expanding Deadly Conflict. Africa Report No. 252.

are derived from FAO Statistics and measure the percentage of national territory under pastureland and forest. On average, permanent pastures account for 30.7% of land area in Sub-Saharan Africa and have remained relatively stable (-0.5% change over the period). In contrast, forest coverage declined by 6.7%, with an average of 32.5% of land classified as forest.

Previous research shows that politically excluded or economically marginalized ethnic groups are more likely to rebel (e.g., Cederman, Gleditsch and Buhaug, 2013). This evidence suggests that inequality can lead to group mobilization and violence and recognizes the importance of organizations in fostering collective action (Vogt, Gleditsch & Cederman, 2021). Furthermore, other scholars emphasize that ethnic grievances are shaped not only by structural inequalities but also by strategic factors, such as a group's organizational strength and bargaining power (Jenne, Saideman & Lowe, 2007). In contexts where information is limited and distrust is high, ethnic grievances can further increase tensions, especially when excluded groups compete for access to the same resources, leading to outbreaks of violence at the local level. For this reason, in our analysis, we also include the size of the discriminated population as a control variable, since larger groups may be more able to mobilize and have a greater impact on local conflict dynamics. We rely on the Ethnic Power Relations Dataset (Vogt et al., 2015) that provides annual data on politically relevant ethnic groups, their relative sizes as a share of the total population, and codes their access to state power to measure ethnic discrimination. On average, 3.25% of the population is classified as politically discriminated against during the observation period. However, this proportion is substantially higher in certain contexts: for example, it reaches approximately 18% in countries such as Kenya where some ethnic groups maintain a systematic dominance over bureaucracy and political elites, often to the detriment of more vulnerable populations, such as pastoralist communities and forest tribes[9].

We also control for GDP per capita in our analysis, as economic capacity is one of the most robust correlates of conflict occurrence. Lower levels of economic development often coincide with limited state capacity, higher poverty rates, and lack of economic opportunities for stable livelihoods, all of which can exacerbate grievances and increase the appeal of participation in violence [10]. In contrast, higher income levels tend to correlate with improved living standards and lower incentives to engage in conflict. Therefore, including GDP per capita helps isolate the specific effects of resource-related variables on conflict by accounting for the broader economic context.

In addition, we control the share of the total population living in rural areas and, through an additional dichotomous variable, whether this exceeds the 75% threshold found in the distribution to signal a predominantly rural country. As population density increases, competition for limited resources such as land, water, and food intensifies. This scarcity can exacerbate tensions between different groups, leading to conflict. For example, in northern Nigeria, the increase in violence has been attributed to long-standing disputes over land and resources between nomadic herders and sedentary farmers, a situation worsened by climate change and population pressures. By controlling for rural population, we can better isolate the effects of the overall population size on the risk of communal violence, leading to more accurate and nuanced findings.

Lastly, we include a control for past incidents of communal violence. This temporal lag accounts for the well-documented recursive nature of violence, where previous episodes increase the likelihood of future ones. Recurrence may result from unresolved grievances, cycles of revenge, weakened trust among groups, or the institutionalization of violent strategies as a means of conflict resolution. The persistence of violence over time is a robust finding in the conflict literature, supported by studies in different contexts (see, for example, Walter, 2004).

[9] See, for example, the Minority and Indigenous Trend 2023 report available at https://minorityrights.org/resources/?content_type%5B0%5D=flagship-publications

[10] This correlation has been well established since the seminal studies by Collier and Hoeffler, 2004.

5 *Empirical analysis*

The analysis begins by exploring the determinants of the likelihood of communal violence as specified in the baseline model. In this initial stage, sets of control variables are progressively introduced. This stepwise inclusion comes at the cost of sample size, as the availability of data varies across specifications: while Model 1 benefits from full temporal coverage up to 2024, Models 4.1 and 4.2 can only be estimated through 2021[11].

The results consistently point to a negative and statistically significant association between the degree of legal transparency and predictable enforceability of the laws and the likelihood of communal violence (Table 1). This stability across models suggests that the finding is not driven by model specification or changes in sample coverage, but rather reflects a robust underlying relationship.

Economic development, measured as (ln) GDP per capita, is negatively associated with communal violence in the baseline models. However, this effect loses statistical significance once additional controls are introduced, in particular land-use variables that in our framework capture livelihood structures. By contrast, the share of rural population emerges as a strong and stable predictor of violence, while the dummy for predominantly rural countries remains imprecisely estimated throughout. Land-use variables provide more mixed evidence: pasture land shows no consistent effect, whereas forest land is negatively associated with violence in a substantively large and statistically significant way, which may reflect the protective role of forest coverage in reducing competition over agricultural resources. Finally, the share of discriminated population and the lagged occurrence of past communal violence both have positive and significant coefficients, in line with theoretical expectations regarding the role of social exclusion and conflict persistence.

Table 1: Likelihood of events of communal violence

	(1)	(2.1)	(2.2)	(3.1)	(3.2)	(4.1)	(4.2)
Law transp. & predict.	-0.376***	-0.356***	-0.418***	-0.371***	-0.438***	-0.339***	-0.367***
	(0.121)	(0.127)	(0.111)	(0.144)	(0.120)	(0.113)	(0.095)
(ln) GDPpc	-0.554*	-0.596**	-0.563*	-0.544*	-0.485	-0.243	-0.234
	(0.330)	(0.294)	(0.308)	(0.316)	(0.326)	(0.250)	(0.225)
(ln) Rural pop.	1.162***	1.187***	1.063***	1.279***	1.123***	0.939***	0.853***
	(0.265)	(0.263)	(0.234)	(0.282)	(0.229)	(0.197)	(0.161)
Predominantly rural	-0.216	-0.198	-0.401	-0.227	-0.442	-0.094	-0.352
	(0.350)	(0.331)	(0.340)	(0.342)	(0.355)	(0.295)	(0.295)
Pasture land (%)		2.444		2.819		1.069	
		(1.640)		(2.048)		(1.223)	
Forest land (%)			-3.652***		-3.900***		-3.068***
			(1.258)		(1.309)		(1.011)
Discriminated pop. (%)				1.492*	1.504**	1.346**	1.300**
				(0.784)	(0.714)	(0.608)	(0.536)
Past communal violence						1.030***	1.058***
						(0.205)	(0.201)
Observations	1418	1418	1418	1336	1336	1336	1336
Number of Groups	43	43	43	43	43	43	43
AIC	653.229	652.085	646.701	608.075	602.343	567.690	558.013
BIC	684.771	688.884	683.500	649.654	643.923	614.467	604.790

Significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Note: Standard errors clustered at country level in parentheses. All variables are temporally lagged one year. Since data on the presence of discriminated groups are available only up to 2021, the number of observations decreases accordingly from Model 3.1 onward, while the number of groups is preserved.

When climate variability is introduced to assess its role on the effect of institutional quality, the results reveal substantively important implications for the interpretation of the institutional–conflict relationship. Table

[11] As a robustness check, we re-estimate the baseline model on the restricted sample, which has a shorter temporal coverage for all countries due to data availability. The results remain fully consistent with those obtained from the full sample, indicating that the main findings are not driven by selective data loss. Detailed results are reported in Appendix B

2 reports estimates of the likelihood of communal violence events while explicitly accounting for dry and wet climatic conditions, as measured by SPEI, and their interaction with legal transparency and predictable enforcement.

Across all specifications, legal transparency and predictability exhibit a robust and statistically significant negative association with communal violence, confirming their conflict-mitigating role under both drier and wetter conditions. At the same time, climate variability does not display a statistically significant direct effect on the likelihood of communal violence, regardless of whether conditions are drier or wetter. This suggests that climatic fluctuations, in isolation, are not sufficient to explain variations in conflict probability. This result is consistent with (Koubi, 2019).

However, the interaction terms indicate that the relevance of climate variability emerges through its interplay with institutional quality. Under drier conditions, the interaction between legal transparency and climate variability is positive and statistically significant, indicating that the violence-reducing effect of strong legal institutions weakens as drought intensity increases. This pattern implies that climatic stress may erode the capacity of even relatively transparent and predictable legal systems to prevent communal conflict by amplifying resource competition and grievances that overwhelm institutional conflict-resolution mechanisms.

By contrast, no statistically significant interaction emerges under wetter conditions. Taken together, these results highlight an asymmetric and indirect role of climate variability: rather than acting as an independent driver of communal violence, climate variability becomes consequential insofar as it interacts with institutional mediators that shape societies' ability to manage potential conflicts.Finally, it is worth noting that, across all model specifications, controlling for forest share rather than pasture share systematically improves goodness-of-fit measures, suggesting that forest cover captures relevant variation in land use more closely related to the likelihood of communal violence.

Table 2: Likelihood of Communal Violence Events, Controlling for Climate Variability (SPEI)

	Drier conditions		Wetter conditions	
	(5.1)	(5.2)	(6.1)	(6.2)
Law transp. & predict.	-0.474***	-0.491***	-0.339***	-0.368***
	(0.129)	(0.119)	(0.112)	(0.090)
SPEI_dry	-0.004	0.009		
	(0.117)	(0.118)		
Law transp. & predict.*SPEI_dry	0.198**	0.185**		
	(0.083)	(0.082)		
SPEI_wet			-0.267	-0.279
			(0.205)	(0.203)
Law transp. & predict.*SPEI_wet			-0.018	-0.015
			(0.161)	(0.155)
(ln) GDPpc	-0.217	-0.211	-0.222	-0.221
	(0.248)	(0.225)	(0.244)	(0.219)
(ln) Rural pop.	0.942***	0.859***	0.960***	0.877***
	(0.197)	(0.161)	(0.199)	(0.163)
Predominantly rural	-0.081	-0.334	-0.067	-0.327
	(0.296)	(0.293)	(0.303)	(0.303)
Pasture land (%)	1.025		1.015	
	(1.216)		(1.223)	
Forest land (%)		-2.980***		-3.123***
		(1.010)		(1.025)
Discriminated pop. (%)	1.351**	1.310**	1.294**	1.252**
	(0.607)	(0.539)	(0.593)	(0.522)
Past communal violence	1.041***	1.067***	1.025***	1.050***
	(0.199)	(0.196)	(0.201)	(0.197)
Observations	1336	1336	1336	1336
Number of Groups	43	43	43	43
AIC	566.221	557.103	569.290	559.339
BIC	623.393	614.275	626.461	616.510

Significance levels: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$. Standard errors clustered at country level in parentheses. All variables are temporally lagged one year.

Taken together, the results in Table 2 reinforce the central role of legal transparency and predictability in shaping the risk of communal violence, while showing that their effectiveness is conditionally weakened under adverse climatic conditions. While these findings are consistent with the proposed theoretical framework, they are based on associational evidence and should be interpreted with caution. In particular, given the strong and persistent relationship between institutional quality and communal violence, it is important to examine whether the estimated effects may be affected by factors that could give rise to endogeneity, such as reverse causality or unobserved, time-invariant determinants jointly influencing institutions and conflict outcomes.

5.1 Robustness Checks for Institutional Endogeneity

We assess the robustness of our main findings to potential endogeneity in institutional quality using the two complementary approaches outlined in Section 3: a Mundlak-style correlated random effects specification to account for unobserved, time-invariant country-specific factors, followed by a Control Function analysis to examine possible reverse causality.

In a first stage, we include country-specific time averages of all covariates as additional regressors in a Mundlak-style correlated random effects specification. This accounts for potential unobserved, time-invariant heterogeneity while retaining all countries in the sample.

Table 3 reports Wald tests for the joint significance of these mean terms. In most model specifications, the tests fail to reject the null, indicating that the Mundlak terms are generally insignificant, supporting the robustness of the random-effects probit estimates reported in Table 2. An exception is the model controlling for forest share and wetter-than-average conditions (Model 6.2), where the Wald test rejects the null ($\chi^2 = 21.27$,

$p = 0.0065$). Inspection of the coefficient estimates suggests a marginally significant negative relationship between institutional quality and communal conflict under these conditions. Thus, the effect of institutions on communal conflict is confirmed, although limited; however, the positive interaction with wetter-than-normal conditions reduces this mitigating effect as the climate becomes wetter and may even reverse it under extreme wet conditions. This pattern is observed only in this specification and, as previously noted, could suggest that controlling for forest share captures structural characteristics more strongly associated with the occurrence of communal conflict. The results are reported in detail in Appendix C.

Table 3: Wald tests for Mundlak covariate means in communal conflict models

Climate conditions	**Model**	χ^2	**Prob >** χ^2
Drier conditions	5.1	12.78	0.3163
	5.2	12.78	0.1195
Wetter conditions	6.1	8.58	0.3792
	6.2	21.27	0.0065

Note: Wald tests evaluate the joint significance of the country-specific time averages of all covariates included in the Mundlak-style correlated random effects specifications.

In a second stage, we further explore a potential source of endogeneity in the institutional variable using a Control Function approach[12]. In this procedure, we first regress the variable of interest on the instruments and other control variables, and then include the residuals from this first-stage regression as an additional regressor in the main model. This allows us to assess whether reverse causality—from conflict back to institutional quality—may be relevant, based on the statistical significance of the residuals. As part of the approach, we use as an instrument the average value of the main regressor measured during the first five years of a country's independence. This period corresponds to the formative stage of state-building, during which the basic legal and institutional architecture of the new polity is established.

Since the instrument is constructed as a country-specific average over the first five years following independence, it varies across polities while remaining effectively time-invariant throughout the sample period. This is supported by the observation that roughly 83% of the total variance in the instrument is between countries, with only 17% occurring within countries over time, implying that cross-country differences primarily drive the variation in the independent variable. Appendix A provides data on the countries included in the study, their year of independence, and the values of the instrument. A cluster-robust Wald χ^2 from the first-stage regression, which serves as a close approximation to the conventional Cragg–Donald / Staiger–Stock F, indicates that the instrument is strongly correlated with the main regressor, achieving values well above conventional threshold (Stock & Yogo, 2005) when controlling for both pasture share ($\chi^2 = 25.508$, p=0.000) and forest share ($\chi^2 = 19.89$, p=0.000).

Table 4 presents the results of the panel probit estimations for the likelihood of communal violence, where we apply a Control Function approach using a deep lag of institutional quality as an instrument to explore the potential role of reverse causality. These models build on the specifications presented in Table 2, allowing a

[12]As a preliminary check, we also estimated the reversed specification, testing whether previous occurrences of communal conflict could predict current levels of law transparency and predictable enforcement. The coefficient was not statistically significant (p-value = 0.428), ruling out this potential reverse-causality concern. Nevertheless, to further explore this possibility, we apply a Control Function approach to assess whether residual correlation between conflict and institutional quality could influence the estimated effects.

direct comparison while accounting for the first-stage residuals. All other covariates exhibit patterns consistent with our baseline models, confirming the robustness of previous findings[13].

Focusing on the interaction between legal transparency and enforceability and climate variability, the results indicate that under drier conditions, the interaction remains positive and statistically significant, suggesting that the violence-reducing effect of stronger legal institutions weakens as drought intensity increases. By contrast, no significant interaction emerges under wetter conditions, consistent with the asymmetric role of climate highlighted in earlier results.

The first-stage residuals included as part of the Control Function approach are not statistically significant in any specification, providing no indication of substantial feedback from conflict to institutions. Formal Wald tests fail to reject the null hypothesis that the residual coefficient equals zero in specifications accounting for both pasture share and forest share (Table 5). Across both drier and wetter conditions, the residuals are not statistically different from zero (all p-values > 0.49), suggesting that our main estimates are robust to potential reverse causality.

Taken together, these results suggest that the main findings remain stable when accounting for unobserved heterogeneity and when considering alternative specifications designed to address concerns related to reverse causality.

Table 4: Control Function approach for likelihood of communal violence

	Drier conditions		Wetter conditions	
	CF(5.1)	CF(5.2)	CF(6.1)	CF(6.2)
Law transp. & predict.	-0.676**	-0.679**	-0.537	-0.550*
	(0.332)	(0.324)	(0.335)	(0.310)
SPEI_dry	-0.005	0.008		
	(0.118)	(0.120)		
Law transp. & predict. × SPEI_dry	0.197**	0.184**		
	(0.083)	(0.084)		
SPEI_wet			-0.260	-0.269
			(0.202)	(0.200)
Law transp. & predict. × SPEI_wet			-0.034	-0.025
			(0.156)	(0.152)
1st stage residulas	0.256	0.252	0.236	0.229
	(0.379)	(0.386)	(0.343)	(0.351)
(ln) GDPpc	-0.210	-0.200	-0.217	-0.212
	(0.281)	(0.259)	(0.278)	(0.255)
(ln) Rural pop.	0.986***	0.861***	1.006***	0.877***
	(0.234)	(0.179)	(0.241)	(0.185)
Predominantly rural	-0.075	-0.362	-0.062	-0.360
	(0.341)	(0.341)	(0.351)	(0.356)
Pasture share (%)	1.207		1.198	
	(1.256)		(1.281)	
Forest (%)		-3.395***		-3.529***
		(1.262)		(1.298)
Discriminated pop. (%)	1.151**	1.144**	1.096*	1.093**
	(0.568)	(0.499)	(0.565)	(0.485)
Past communal violence	0.999***	1.029***	0.981***	1.011***
	(0.196)	(0.197)	(0.200)	(0.201)
Observations	1280	1280	1280	1280
Number of Groups	40	40	40	40
AIC	562.610	553.173	565.554	555.318
BIC	624.466	615.028	627.409	617.174

Significance levels: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$. Standard errors clustered at country level in parentheses. All variables are lagged by one year, except for the institutional quality indicator, which is a deep lag measured as the average over the first five years following a country's independence. Results reported in the table correspond to second-stage estimates from the Control Function approach, obtained using panel probit models.

[13] See Appendix C for the first-stage Control Function results.

Table 5: Wald tests for control function residuals in communal conflict models

Climate conditions	Model	Residual	χ^2	Prob > χ^2
Drier conditions	5.1	resid_p	0.46	0.4998
	5.2	resid_f	0.47	0.4919
Wetter conditions	6.1	resid_p	0.43	0.5139
	6.2	resid_f	0.42	0.5145

Note: `resid_p` is the residual from the control function including `pasture_share`, whereas `resid_f` is the residual from the control function including `forest_share`. Wald tests test whether each residual is significantly different from zero.

Two limitations should be acknowledged. First, despite extensive robustness checks, the empirical strategy does not allow for a fully causal interpretation of the estimated effects, as institutional quality may still be correlated with time-varying unobserved factors. Second, the control-function approach using deep institutional lags is best interpreted as a diagnostic tool for reverse causality rather than a source of causal identification.

Within these limits, however, the robustness exercises performed in this section are informative about the nature of the relationship under study. In particular, the stability of the estimated coefficients and interaction terms across alternative specifications, correlated random effects, and control-function models suggests that the main results are not driven by mechanical feedback from conflict to institutions nor by time-invariant country-specific heterogeneity. Rather than establishing causality, these checks strengthen the interpretation of the estimated relationships as systematic and structurally meaningful associations, which are consistent with the proposed theoretical mechanisms linking institutional quality, climatic stress, and communal violence.

5.2 Institutional Quality, Climate Variability, and Communal Conflict

Returning to the results presented in Table 2, this section discusses the role of institutional quality and its interaction with climate variability in explaining communal conflict.

While the estimated coefficients indicate the presence of a statistically significant interaction between institutional quality and drought severity, their substantive interpretation in a non-linear model is not straightforward. To facilitate interpretation, we therefore compute and report average marginal effects of institutional quality on the probability of communal conflict across different levels of drought severity.

Figure 5 reports the average marginal effect of institutional quality on the probability of communal conflict across increasing levels of drought severity (`SPEI_dry`). Under conditions of low or negligible drought severity, a one-unit increase in institutional quality is associated with a reduction in the probability of communal conflict of approximately 5–6 percentage points. This mitigating effect weakens monotonically as drought severity increases, with the estimated average marginal effect approaching zero at SPEI values of approximately 2.5–2.7 in both specifications. Note that the original SPEI index is expressed in absolute value, such that higher values correspond to greater drought severity. This pattern is highly consistent across alternative model specifications controlling for land-use composition. The average marginal effects derived from models including pasture share and forest share follow nearly identical trajectories across the range of drought severity, differing only marginally in magnitude. These small differences do not affect the substantive interpretation of the results, indicating that the erosion of the conflict-mitigating role of institutional quality under increasing drought severity is robust to alternative land-use controls.

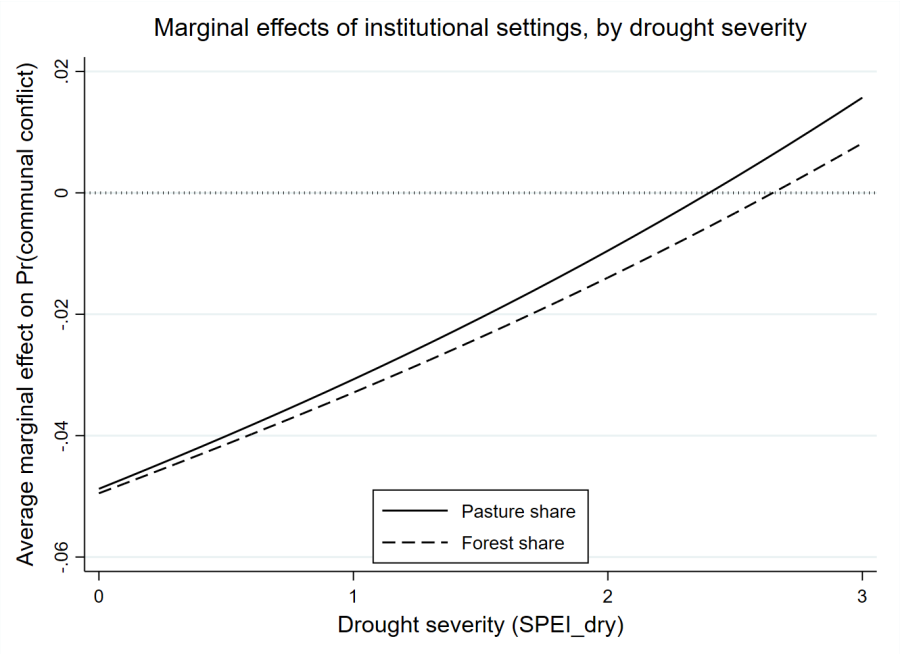

Figure 5: Average marginal effect of institutional quality on the probability of communal conflict across increasing levels of drought severity (SPEI). The figure reports average marginal effects from panel probit models with country random effects. Solid and dashed lines correspond to alternative model specifications controlling for pasture (Mod. 5.1) and forest land share (Mod. 6.1), respectively.

6 *Conclusions*

This paper examines the relationship between land-related institutional settings and communal violence in Sub-Saharan Africa, with particular attention to the conditional role of climate shocks. Using a country–year panel covering the period 1990–2024 and focusing on communal violence events, the analysis documents a consistent negative association between greater legal transparency and predictable enforcement of laws and the likelihood of communal violence. This relationship remains robust across model specifications and alternative controls for land use, and is further supported by robustness checks accounting for potential reverse causality and unobserved heterogeneity.

The results highlight that the formal quality of legal institutions—captured by transparency, coherence, and predictability of enforcement—is systematically associated with a lower likelihood of communal violence. Rather than focusing on specific land tenure arrangements, which are difficult to observe in a comparable manner across countries and over time (Deininger et al., 2012), the analysis relies on an institutional dimension that also captures how access to land and land-related resources is regulated, adjudicated, and enforced through the legal system. In contexts where land constitutes a primary livelihood asset and competition over common-pool resources is pervasive, clear and consistently enforced rules are associated with lower uncertainty surrounding land access and use, a pattern consistent with reduced incentives for violent mobilization (Acemoglu et al., 2001; North, 1990). This dimension remains a strong and robust correlate of lower conflict risk even after accounting for economic development, rural population pressure, land-use patterns, ethnic discrimination, and conflict persistence.

At the same time, the analysis shows that the conflict-mitigating effect of institutional quality is not uniform across environmental conditions. Climate variability does not emerge as an independent driver of communal violence; rather, its relevance is revealed through conditional effects on institutions (McGuirk & Nunn, 2020). Under drier conditions, increases in drought severity systematically weaken the association between legal trans-

parency and predictability and lower levels of communal violence. Average marginal effects indicate that while improvements in institutional quality are associated with substantial reductions in conflict risk under normal climatic conditions, this protective association progressively erodes as drought intensity rises, approaching zero under severe drought. By contrast, no comparable interaction is observed under wetter conditions. This evidence is consistent with a broader literature emphasizing that environmental stress affects conflict risk primarily through economic and institutional channels rather than through direct effects (Koubi, 2019; Miguel et al., 2004).

These findings support the interpretation of climate shocks—particularly slow-onset events such as droughts — as a threat multiplier rather than a direct cause of communal violence (Froese & Schilling, 2019). Drought-induced environmental stress appears to amplify underlying vulnerabilities, likely by intensifying competition over land and natural resources, thereby straining institutional mechanisms that would otherwise mitigate conflict. Even relatively strong and predictable legal systems may be less effective when resource scarcity becomes acute and grievances accumulate faster than institutions can manage them.

The paper provides novel cross-country evidence on how land-related institutional settings and environmental conditions jointly shape the risk of communal violence, offering a coherent empirical framework that can inform both future causal research and policy discussions in land-dependent and climate-vulnerable contexts. From a policy perspective, the results underscore the importance of complementing institutional strengthening with targeted climate-adaptation strategies, particularly in drought-prone settings. While improving legal transparency and predictability remains a cornerstone of violence prevention, such reforms may be insufficient in isolation when societies face acute climatic stress. In this respect, policies aimed at reducing vulnerability to drought—such as investments in water management, livelihood diversification, and early warning systems—can help preserve the conflict-mitigating role of institutions under adverse environmental conditions. Taken together, the findings suggest that institutional arrangements and environmental conditions should be considered jointly when assessing the risk of communal violence in land-dependent societies.

A Summary Statistics

This section provides an overview of the variables employed in the study. Table A1 reports descriptive statistics, including the mean, standard deviation, minimum and maximum values, and data sources. Table A2 lists the countries in the analysis, indicating whether they experienced at least one episode of communal violence and the total number of such events. It also reports the average degree of transparency and predictable enforceability of laws over the sample period, as well as each country's year of independence and the average value of the same institutional indicator in the five years following independence, used as an instrument in the Control Function analysis. Finally, Table A3 presents the correlation matrix, illustrating the pairwise relationships among the variables.

Table A1: Descriptive statistics of the variables

Variable	Obs	Mean	Std. Dev.	Min	Max	Source
Communal conflict	1,481	0.1985	0.3990	0	1	UCDP-GED v25_1
Law transp. & predict.	1,481	0.2007	1.2410	-2.748	2.820	V-Dem v15
CF_Law transp. & predict.	1,400	-0.3194	0.9267	-2.623	1.292	V-Dem v15
SPEI_dry	1,438	0.5428	0.6230	0	3.0459	ESG Data Portal (WB)
SPEI_wet	1,438	0.2280	0.4217	0	3.0738	ESG Data Portal (WB)
Pasture land (%)	1,438	0.3076	0.1845	0.0010	0.6919	Faostat
Forest land (%)	1,438	0.3246	0.2486	0.0029	0.9623	Faostat
(ln) GDPpc	1,459	6.9124	0.8536	5.2399	9.4764	WDI
(ln) Rural pop.	1,481	15.577	1.3610	12.3046	18.466	WDI
Predominantly rural	1,481	0.3336	0.4716	0	1	calculated from WDI
Discriminated pop. (%)	1,352	0.0326	0.1017	0	0.8600	Ethnic Power Relations Dataset

Table A2: List of Countries, Independence Dates, and Instrument Values

Country	Comm. confl.	Total events	Law transp. & predict.	Independence year	CF_Law transp. & predict.
Angola	no	0	-0.6154	1975	-1.0480
Benin	no	0	1.1321	1960	-0.4156
Botswana	no	0	1.5889	1966	1.2230
Burkina Faso	no	0	1.7052	1960	0.9440
Burundi	yes	256	-0.3634	1962	-1.7914
Cameroon	yes	15	0.0344	1960	-1.6170
Central African Republic	yes	111	-0.4673	1960	0.4080
Chad	yes	45	-1.1030	1960	-1.1150
Dem. Rep. of the Congo	yes	414	-1.2460	1960	-2.2184
Equatorial Guinea	no	0	-1.3919	1968	-2.6232
Eritrea	no	0	-2.1015	1993	.
Eswatini (Swaziland)	no	0	-1.0718	1968	-0.5900
Ethiopia	yes	402	0.0125	*1960*	-0.2032
Gabon	no	0	1.2210	1960	-0.4540
Ghana	yes	55	1.8369	1957	0.5474
Guinea	yes	9	-0.5651	1958	-1.1998
Guinea-Bissau	no	0	0.2882	1973	-0.8558
Ivory Coast	yes	42	0.8473	1960	0.4670
Kenya	yes	565	0.6442	1963	-0.8594
Lesotho	no	0	1.5111	1966	0.1826
Liberia	no	0	-0.3281	*1960*	-0.4180
Madagascar	no	0	0.0296	1960	-0.5080
Malawi	no	0	1.4041	1964	-1.1316
Mali	yes	39	0.9535	1960	-0.1570
Mauritania	no	0	-0.0490	1960	-0.0210
Mozambique	no	0	0.6963	1975	0.3722
Namibia	no	0	1.8996	1990	.
Niger	yes	13	2.3225	1960	0.9740
Nigeria	yes	1607	-0.7945	1960	-0.3660
Republic of the Congo	no	0	-0.5796	1960	-0.4850
Rwanda	yes	5	0.7428	1962	0.7954
Senegal	no	0	1.9340	1960	1.1152
Sierra Leone	no	0	-0.3481	1961	0.5010
Somalia	yes	219	-1.6031	1960	-0.9612
South Africa	yes	3	1.2564	1961	-0.5960
South Sudan	yes	211	-1.7641	2011	.
Sudan	yes	478	-1.5715	1956	-0.7586
Tanzania	no	0	1.2927	1961	0.8360
The Gambia	no	0	-0.0399	1965	1.2920
Togo	yes	1	-0.4141	1960	-1.1952
Uganda	yes	93	0.3787	1962	-0.7968
Zambia	no	0	0.7517	1964	0.1380
Zimbabwe	no	0	-0.8120	1980	-0.1902

Notes: *Law transp.& predict.* reports the average value over the entire observation period (1990–2024). The last column (*CF_Law transp.& predict.*) reports the average value of this institutional dimension over the first five years following independence.

Two cautionary notes. First, Eritrea and South Sudan gained independence as secessions from existing African states, while Namibia obtained independence from a former African-administered mandate. In all three cases, independence occurred very recently and falls within the observation period. These features weaken the rationale for using the post-independence average of law transparency and predictable enforceability as a valid instrument in the years immediately following independence. Consequently, these countries are excluded from the Control Function approach (see Table 4).

Second, for Ethiopia, which experienced a brief Italian occupation but was never formally colonized, and for Liberia, which was never colonized, we define a conventional post-independence period of 1960–1964. This choice allows us to compute average values of institutional measures that are comparable to the first five years following independence in other African countries, ensuring consistency in our Control Function approach.

Table A3: Correlation matrix of the variables

Variable	Comm. confl.	LawTransp. &Predict.	CF_LawTransp. &Predict.	SPEI_dry	SPEI_wet	Pasture	Forest	(ln)GDPpc	(ln)RuralPop.	Pred.Rural	Discr.Pop	Comm. Confl._lag
Comm. confl.	1.0000											
LawTransp.&Predict.	-0.2290	1.0000										
CF_LawTransp.&Predict.	-0.1759	0.5476	1.0000									
SPEI_dry	-0.0783	0.0068	0.0357	1.0000								
SPEI_wet	0.0067	0.0485	-0.0336	-0.4714	1.0000							
Pasture	0.0593	-0.0712	0.1604	0.0267	-0.0217	1.0000						
Forest	-0.2494	-0.0883	-0.1916	-0.0092	-0.0183	-0.5403	1.0000					
(ln) GDPpc	-0.1324	0.1196	-0.0317	-0.0004	0.0114	0.1225	0.2229	1.0000				
(ln) Rural pop.	0.4737	0.0392	-0.0310	-0.0576	0.0573	0.0745	-0.3514	-0.3643	1.0000			
Pred. rural	0.0895	0.1433	0.0264	-0.0036	0.1179	0.1270	-0.4283	-0.3647	0.2392	1.0000		
Discr. pop	0.1714	-0.1519	-0.0843	-0.0215	-0.0072	0.0006	-0.0437	0.0081	0.1357	0.1103	1.0000	
Comm. confl_lag	0.7370	-0.2262	-0.1744	-0.0587	-0.0120	0.0609	-0.2506	-0.1350	0.4734	0.0887	0.1685	1.0000

B Additional estimations

To assess whether the reduction in sample size induced by the inclusion of the additional covariate `Discriminated pop(%)`—which primarily reflects its shorter temporal coverage—affects the baseline results, we re-estimate the baseline specification on the temporally truncated sample of 1,336 observations. The estimated coefficients remain fully consistent with those obtained using the full panel, indicating that the main findings are not driven by the shorter time dimension induced by data availability constraints (Table B1). The reduction in sample size is due to the uniform exclusion of the last years for all countries and a few missing observations, ensuring that the main findings are not driven by selective data loss.

Table B1: Baseline Model Estimates on the Restricted Sample

	(1)	(2.1)	(2.2)	(2.1.1)	(2.2.1)
Law transp. & predict.	-0.430***	-0.408***	-0.470***	-0.368***	-0.389***
	(0.132)	(0.139)	(0.118)	(0.109)	(0.095)
(ln) GDPpc	-0.570	-0.633**	-0.563*	-0.302	-0.271
	(0.358)	(0.310)	(0.331)	(0.248)	(0.219)
(ln) Rural pop.	1.214***	1.253***	1.104***	0.926***	0.844***
	(0.279)	(0.281)	(0.233)	(0.197)	(0.158)
Predominantly rural	-0.268	-0.252	-0.448	-0.092	-0.334
	(0.362)	(0.340)	(0.349)	(0.293)	(0.285)
Pasture land (%)		2.949		1.174	
		(1.965)		(1.206)	
Forest land (%)			-3.868***		-3.048***
			(1.289)		(0.985)
Past communal violence				1.039***	1.073***
				(0.209)	(0.205)
Observations	1336	1336	1336	1336	1336
Number of Groups	43	43	43	43	43
AIC	613.991	612.013	606.935	570.748	561.159
BIC	645.176	648.395	643.317	612.327	602.738

Significance levels: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$. Standard errors clustered at country level in parentheses. All variables are temporally lagged one year.

As an additional robustness check, we re-estimated the reference models using a random-effects logit specification instead of the baseline random-effects probit. This allows us to assess whether the results are sensitive to the choice of the link function and distributional assumptions underlying the binary outcome model. The corresponding estimates are reported in Table B2

Table B2: Panel logit estimations for likelihood of communal violence

	(5.1)	(5.2)	(6.1)	(6.2)
Law transp. & predict.	-0.909***	-0.942***	-0.636***	-0.688***
	(0.237)	(0.219)	(0.206)	(0.167)
SPEI_dry	-0.047	-0.014		
	(0.214)	(0.218)		
Law transp. & predict. × SPEI_dry	0.390**	0.370**		
	(0.156)	(0.154)		
SPEI_wet			-0.434	-0.460
			(0.379)	(0.376)
Law transp. & predict. × SPEI_wet			-0.100	-0.091
			(0.308)	(0.302)
(ln) GDPpc	-0.571	-0.538	-0.529	-0.514
	(0.456)	(0.419)	(0.449)	(0.406)
(ln) Rural population	1.857***	1.672***	1.892***	1.699***
	(0.380)	(0.303)	(0.385)	(0.306)
Predominantly rural	-0.339	-0.811	-0.239	-0.739
	(0.539)	(0.509)	(0.562)	(0.544)
Pasture share (%)	1.968		1.937	
	(2.297)		(2.325)	
Forest share (%)		-5.794***		-6.022***
		(1.914)		(1.941)
Discriminated population (%)	2.452**	2.389**	2.328**	2.260**
	(1.118)	(1.011)	(1.084)	(0.974)
Past communal violence	1.729***	1.786***	1.705***	1.753***
	(0.348)	(0.338)	(0.352)	(0.342)
Observations	1336	1336	1336	1336
Number of Groups	43	43	43	43
AIC	566.250	557.314	570.251	560.556
BIC	623.422	614.485	627.423	617.728

Significance levels: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$. Standard errors clustered at country level in parentheses. All variables are temporally lagged one year.

C Robustness Checks for Institutional Endogeneity: additional results

This section reports additional results related to the robustness checks discussed in the main text (Section 5.1), including i) the Mundlak-style correlated random effects specifications used to account for unobserved heterogeneity and ii) the first-stage estimates of the Control Function approach.

C.1 Checks for unobserved heterogeneity

The Mundlak-corrected random-effects probit models serves as a robustness check of the random-effects specifications as reported in Table 2. Among the estimated models, only Model 6.2—combining institutional quality with wetter-than-average conditions while controlling for forest share—rejects the null hypothesis that the Mundlak terms are jointly insignificant, as indicated by the Wald test (Table 3). The following figure provides a graphical representation of the coefficients and intervals obtained from the Mundlak correction for Model 6.2.

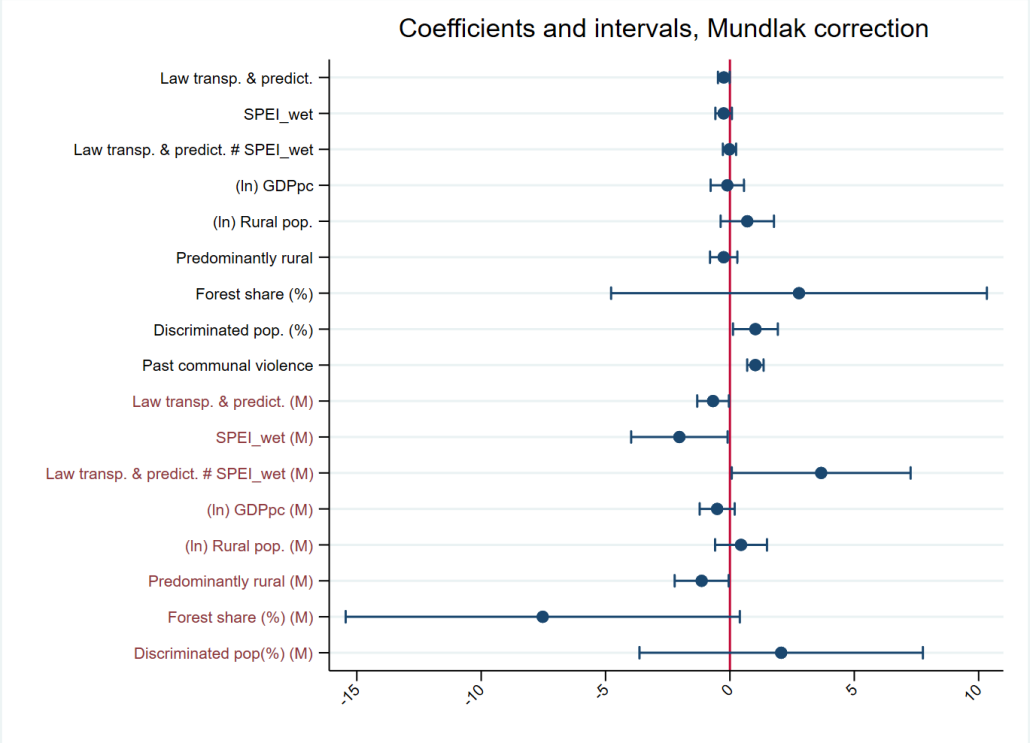

Figure C1: Results for Model 6.2 (Table 2) estimated using a Mundlak-corrected random-effects probit specification, which accounts for unobserved time-invariant heterogeneity by including country-specific time averages of all covariates.

In this Mundlak-corrected random-effects probit model, the joint effect of law transparency and predictability and wetter climate conditions appears suggestive. The main coefficient of the institutional parameter is significantly negative, consistent with previous results showing that land-related institutional settings tend to mitigate the likelihood of communal violence. Importantly, unlike in the main specifications, the model now reveals a marginally significant role for above-average wet conditions (`SPEI_wet`) when considered together with deviations captured by the Mundlak parameters. These Mundlak terms represent the within-group variation—short-term deviations from each country's average level of law transparency and climate variability—allowing the model to detect context-dependent effects that may have been previously obscured. The

combined effect suggests that the protective influence of law transparency may be somewhat attenuated during unusually wet periods, highlighting a nuanced interplay between governance quality and climatic variability. While the evidence remains suggestive rather than conclusive, it indicates that the effectiveness of institutional mechanisms in mitigating conflict may vary with temporary environmental shocks.

C.2 Checks for reverse causality

To complete the analysis based on the Control Function approach, we now present the first-stage results, which confirm the relevance of the selected instrument.

Table C1: Control Function approach for likelihood of communal violence. First-stage results

	Dep. var.: Law transp. & predict.	
	CF(4.1)	CF(4.2)
CF_Law transp. & predict.	0.701***	0.634***
	(0.138)	(0.143)
(ln) GDPpc	0.234	0.218
	(0.145)	(0.145)
(ln) Rural pop.	0.313*	0.222
	(0.175)	(0.198)
Predominantly rural	0.331**	0.312**
	(0.135)	(0.137)
Pasture land (%)	0.153	
	(1.205)	
Forest land (%)		-1.327
		(1.010)
Discriminated pop. (%)	-0.981***	-0.964***
	(0.338)	(0.340)
Past communal violence	-0.066	-0.061
	(0.072)	(0.072)
Observations	1280	1280
number of groups	40	40
R^2	0.3225	0.2637
F-statistics	25.50	18.74

Significance levels: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$. Standard errors clustered at country level in parentheses. All variables are lagged by one year, except for the institutional quality indicator, which is a deep lag measured as the average over the first five years following a country's independence. Results from the first-stage regression of the Control Function approach.

The first-stage results indicate that the instrument `CF_Law transp.& predict.` is strongly correlated with the potentially endogenous variable, as expected. The coefficient is positive and highly significant in both specifications, and the F-statistics exceed the conventional threshold of F>10 (Stock & Yogo, 2005), confirming that the instrument is sufficiently strong. Overall, these results suggest that the first stage performs well, providing a reliable basis for the subsequent control function estimation.

To further assess the robustness of our findings, we conducted a set of sensitivity analyses focusing on the institutional variable for Ethiopia and Liberia, where the definition of the post-independence period is somewhat arbitrary. Specifically, we vary the temporal window used to compute the institutional indicator for these countries. For Liberia, the indicator remains constant (-0.418) throughout 1955–1965, so modifying the observation window within this range does not affect its value. In contrast, for Ethiopia the indicator varies depending on the selected window. The following table presents the results of the main models (as shown in Table 4) when the values for Ethiopia are modified according to different observation windows. The resulting

coefficients remain very close to the baseline estimates, leaving statistical inference unchanged (Table C2). These results suggest that the findings are robust to the choice of temporal window for these countries.

Table C2: Control Function approach: Inference Stability

Window	CF_Law transp.& predict. (Ethiopia)	Mod. CF(5.1)		Mod. CF(5.2)		Mod. CF(6.1)		Mod. CF(6.2)	
		Coeff.	SE	Coeff.	SE	Coeff.	SE	Coeff.	SE
1960–1964	-0.203	-0.671**	0.331	-0.678**	0.326	-0.532	0.333	-0.549*	0.311
1955–1959	-0.256	-0.676**	0.332	-0.679**	0.324	-0.536	0.335	-0.549*	0.310
1965–1969	-0.168	-0.676**	0.332	-0.680**	0.324	-0.536	0.335	-0.550*	0.310

Significance levels: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$. The second column reports the values of the institutional variable for Ethiopia, depending on the selected observation window used in the main analysis. The remaining columns show the coefficients and standard errors of the variable `Law transp.& predict.` as estimated in the reference models, calculated using the modified values for Ethiopia. The first observation window corresponds to the one used for the results reported in Table 4. All other coefficients retain their original significance levels.

As second sensitivity control, we reduced the sample by excluding Ethiopia and Liberia, and performed the same analysis based on a Control Function approach. The results (Table C3) fully confirm the main findings.

Table C3: Control Function approach for likelihood of communal violence. Sensitivity analysis

	Drier conditions		Wetter conditions	
	CF(5.1.1)	CF(5.2.1)	CF(6.1.1)	CF(6.2.1)
Law transp. & predict.	-0.677**	-0.692**	-0.536	-0.558*
	(0.320)	(0.325)	(0.326)	(0.313)
SPEI_dry	-0.029	-0.017		
	(0.120)	(0.122)		
Law transp. & predict. × SPEI_dry	0.200**	0.187**		
	(0.083)	(0.084)		
SPEI_wet			-0.236	-0.244
			(0.212)	(0.209)
Law transp. & predict. × SPEI_wet			-0.044	-0.037
			(0.157)	(0.154)
1st stage residuals	0.253	0.245	0.253	0.239
	(0.369)	(0.347)	(0.377)	(0.355)
(ln) GDPpc	-0.157	-0.148	-0.172	-0.169
	(0.298)	(0.285)	(0.298)	(0.282)
(ln) Rural pop.	0.934***	0.835***	0.956***	0.855***
	(0.235)	(0.197)	(0.244)	(0.203)
Predominantly rural	-0.076	-0.333	-0.071	-0.338
	(0.344)	(0.345)	(0.354)	(0.360)
Pasture share (%)	0.893		0.904	
	(1.237)		(1.266)	
Forest (%)		-3.391***		-3.538***
		(1.299)		(1.336)
Discriminated pop. (%)	1.106*	1.086**	1.066*	1.052**
	(0.575)	(0.504)	(0.574)	(0.494)
Past communal violence	1.041***	1.068***	1.022***	1.048***
	(0.198)	(0.200)	(0.203)	(0.205)
Observations	1216	1216	1216	1216
Number of groups	38	38	38	38
AIC	537.838	528.910	541.385	531.677
BIC	599.077	590.150	602.625	592.917

Significance levels: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$. Standard errors clustered at country level in parentheses. All variables are lagged by one year, except for the institutional quality indicator, which is a deep lag measured as the average over the first five years following a country's independence. Results reported in the table correspond to second-stage estimates from the Control Function approach, obtained using panel probit models. Ethiopia and Liberia are excluded from the sample.

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