

A Multilevel Bayesian Approach to Climate-Fueled Migration and Conflict

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ABSTRACT

Do climate conditions and extreme events fuel conflict and migration? This question has been widely studied using causal designs that exploit natural variation in climate variables, often analyzed with linear fixed-effects models. Yet in this setting, nonlinear relationships, distributional features of outcomes, and spatial heterogeneity can cause these models to violate core assumptions and yield unreliable inferences. We propose a multilevel Bayesian framework that accommodates such features while retaining identification strategies from natural experiments. We illustrate its potential with a representative analysis from the literature of the effect of temperature anomalies on conflict in Somalia. When outcome distributions suited to event counts are combined with partial pooling across regions, the apparent aggregate climate effect disappears and marked regional heterogeneity emerges, with positive associations in only a few southern regions and negative or uncertain effects elsewhere. Extending pooling across time further improves predictive ability. More broadly, the multilevel Bayesian framework offers a general strategy for strengthening both explanatory and predictive inferences about climate and social outcomes, supporting internal and external validity while efficiently accommodating heterogeneity even with small samples. This methodological bridge between econometric identification strategies and statistical modeling provides a robust foundation for interdisciplinary climate-conflict-migration research.

Keywords: Climate, Conflict, Migration, Multilevel model, Bayesian inference, Causal inference

Introduction

Do climate conditions, and notably climate extremes, fuel or lead to conflict and migration? This question has been addressed by a rapidly growing number of empirical studies over the past two decades, with contributions from a variety of disciplines. In parallel, a large number of systematic reviews have followed, assessing both the diversity of methods used to analyze the statistical relationship between climate and social instability and whether a consensus emerges (see Supplementary Information S.1 online for an overview of reviews published in the past dozen years). Taken together, these reviews find mixed evidence of substantial climate impacts, largely reflecting the variability of specifications and data—such as how the onset of civil war is defined, or which types of displacement are considered^{1–8}. The most recent literature reviews and meta-analyses continue to highlight a lack of consensus^{9,10}, underscoring the persistent methodological and empirical challenges in establishing robust causal relationships. However, those studies within the new climate-economy literature¹¹, which uses reduced-form models exploiting variation in climate variables to identify causal effects, tend to suggest a net effect of climate on conflict^{12–14}.

Research designs that leverage such natural or quasi-experiments are positioned to credibly identify causal effects. Here, *quasi-experiments* refer to observational studies in which quasi-randomization of a variable occurs without researcher intervention, approximating the exogeneity of a randomized experiment, though without full random assignment. Applying such designs to the climate-migration-conflict nexus is of increasing relevance as climate conditions that were historically rare become more frequent and as adaptation becomes a more salient concern. However, this nexus departs substantially from the statistical setting that motivates the widespread linear reduced-form model, in at least two important ways. First, multiple features of the data-generating processes in this context diverge, more strongly than in other contexts, from the assumptions of the linear

model. As a result, the best linear approximation to the conditional expectation function—what the linear reduced-form model provides—may not produce the desired information. Second, while causal inference and prediction can certainly be valuable on their own, we argue that causality questions in this area are often ultimately motivated by prediction; the causal effects of climate on conflict and migration are of increasing practical interest due to expected global changes in the distribution of climate. In effect, causality studies often make implicit or explicit prediction statements, based on their estimates of slope coefficients for climate regressors. Yet with many fixed effects and typically low explained variance, the uncertainty around such predictive statements may be rather large. This motivates turning to approaches that evaluate predictive performance explicitly, while accommodating and estimating heterogeneity in climate effects to support both explanation and prediction. The climate-migration-conflict nexus thereby prompts the use of statistical approaches that integrate both causal identification and predictive validity to provide meaningful insights. While traditional reduced-form linear models with fixed effects remain the standard approach in the causal inference literature, their limitations in capturing nonlinear climate responses and spatial heterogeneity efficiently limit explanatory power and policy relevance.

This paper proposes a multilevel Bayesian framework to analyze climate-fueled social instability. By explicitly modeling heterogeneous regional effects and nonlinear responses while leveraging quasi-experimental identification strategies, this framework offers a way of addressing the aforementioned limitations. We apply it to a dataset representative of the literature and directly compare its results with that from a linear fixed-effects approach. The application illustrates how the latter approach can not only obscure important heterogeneity in climate responses, but also mislead about the aggregate effect. In contrast, the multilevel Bayesian framework improves both the internal validity of causal estimates and out-of-sample predictive ability. In our application to the effect of temperature extremes on the number of conflict events in Somalia, we find that the apparent aggregate climate effect disappears once an outcome distribution suited to event counts is specified and regional pooling is introduced. Only a few southern regions drive positive estimates, whereas others show negative or highly uncertain effects. Moreover, regional partial pooling and then further temporal pooling incrementally improve predictive ability.

The remainder of the paper proceeds as follows. We first describe how specific features of climate-conflict data violate key assumptions of fixed-effects linear models, affecting both causal inference and predictive performance, and why these complementary statistical goals should be jointly considered. Next, we introduce the general multilevel Bayesian framework and describe a simple model for longitudinal data. We then illustrate its potential with the Somalia application and compare the results to those from the fixed-effects linear model. We finally discuss the broader implications of our findings for environmental social science research.

Assumptions and constraints of fixed-effects linear models

Research designs focused on quasi-experiments with adjustments for confounders commonly use a reduced-form linear model to identify an average treatment effect. With longitudinal data, variation in climate variables is leveraged using a multivariate linear regression model, which includes indicator variables or ‘fixed effects’ for the units—e.g., locations—in the sample, as well as temporal fixed effects, to estimate the average effects of climate variables across units and periods. The multivariate linear regression model with fixed effects (MLR-FE) has the following general form:

$$y_{it} = \mathbf{W}'_{it}\beta + \mathbf{X}'_{it}\delta + \phi_i + \psi_t + e_{it}, \quad e_{it} \stackrel{\text{iid}}{\sim} \text{normal}(0, \sigma); \quad i = 1, \dots, n, \quad t = 1, \dots, T \quad (1)$$

where y_{it} is the outcome variable for unit i at time t , \mathbf{W}_{it} is a vector of climate variables considered as treatment (often linear, sometimes including polynomials), \mathbf{X}_{it} is an optional vector of adjustment variables, ϕ_i is a vector of unit fixed effects, and ψ_t is an optional vector of temporal fixed effects. The vector of key parameters of interest β is estimated either by maximum likelihood, with the explicit assumption of normally-distributed errors e_{it} , or by least squares—where the same assumption is implicit for the typical tests of statistical significance. The motivation for using spatial and temporal fixed effects is to adjust for time-invariant and space-invariant unobserved confounders, respectively.

Assuming a strong causal identification strategy, the estimators of the slope parameters are unbiased conditional on the ability to adjust for all confounders. However, the validity, individual interpretation, and assessment of significance of the estimates $\hat{\beta}$ still rely on the modeling assumptions of additivity, linearity, spherical errors, normally-distributed errors, and non-collinear regressors¹⁵. No assumption is ever expected to be met perfectly with real-world data, and exact compliance with all these assumptions is not required for consistency. But in the climate-migration-conflict nexus in particular, the features of the data-generating process often suggest particularly large departures from the assumptions of the MLR-FE model which can become consequential for inference, and include:

- *Nonlinear functional forms.* Migration or conflict outcomes often have nonlinear relationships to environmental conditions, with frequent threshold effects¹⁶. In that context, the simple linear approximation to the conditional expectation function, even if accurate in capturing an average relationship, would fail to capture nonlinearities and

thereby severely limit, if not mislead, the information carried by the slope coefficients, and have weaker predictive ability than a model that captures this behavior.

- *Limited outcome data.* In large samples, OLS estimators remain consistent under non-normal errors. But in this field, outcome data are measures of migration or conflict whose ranges of possible values are often limited—e.g., as they take the form of counts of rare events—producing strongly non-normal conditional distributions. Under such conditions, and especially with relatively small sample sizes, the reliability of conventional significance tests is diminished. This motivates outcome distributions better suited to the nature of the data, and systematic use of diagnostics. Short records across a large number of spatial units also raise the risk of high-influence observations, given the presence of extremes of the often right-skewed climate variables, or of the highly-skewed outcome variables.
- *Correlated climate regressors.* The regression on several climate variables—such as temperature, precipitation, and indices derived from them—limits the simultaneous interpretation of coefficients as individual effects when these variables are highly correlated. In such cases, only their joint effect can be interpreted reliably, while slope estimates risk instability. The literature notes concerns about multicollinearity; some studies address it for example by computing pairwise correlations among climate variables¹⁷ or by explicitly cautioning against interpreting coefficients simultaneously¹⁸.
- *Dependence structures and heterogeneous treatment effects.* Socio-ecological data in the climate-migration-conflict nexus often have a specific dependence structure—whether temporal, spatial or administrative⁵. Longitudinal data have a hierarchical structure, where a lower level is the repeated measure within the group across time and is nested within a higher level which represents the group-level data. So-called "sandwich" estimators, such as those proposed by Conley (1999)¹⁹ and Newey-West (1987)²⁰, are a common way to adjust otherwise misleading estimates of parameter uncertainty for spatial and temporal autocorrelation. However, such adjustment is only valid to the extent that the dependence structure is correctly specified and the sample is sufficiently large. The MLR-FE model also typically fits a separate intercept per group and a homogeneous treatment effect, thereby modeling independent baseline outcome levels across groups, while assuming away similar group-level heterogeneity in the effect of the regressors. This *no-pooling* of intercepts and *full-pooling* of slopes corresponds not to absent or conservative assumptions, but instead to specific restrictive assumptions on the role of hierarchical structures in the data-generating process.

Predictive ability strengthens causal inference

Research questions on the statistical relationships between climate and social outcomes that go beyond merely describing associations can often be classified as belonging to one of three categories:

1. Forward causal inference or "what if" questions, which seek to uncover the effects of causes²¹, e.g., '*What is the average effect of heat waves on migration and conflict?*'
2. Reverse causal inference or "why" questions, regarding causes of effects, e.g., '*What are the causes of the increased rates of interpersonal conflict observed in location X in the last few decades?*'
3. Prediction, e.g., '*How many annual international migrants from region A to region B are expected by mid-century under a global average temperature increase of 2°C above pre-industrial levels?*'

In the heterogeneous body of work concerned with capturing how changes in climate impact conflict and migration, the distinction between these three proximate goals and the focus on different assumptions between disciplines partly explain the variety of statistical approaches. In particular, the divide between explanatory and predictive goals has consequences at every step of the modeling process, from data preparation to the choice of regressors and model selection, while some steps may even virtually disappear, such as model evaluation in causal inference settings²².

In the climate-economy literature, studies generally state the goal of forward causal inference. At the same time, the motivation for prediction is often present—even if implicit—in estimating climate impacts. This is the case in causal inference studies in general: Rubin (1974)²³—the seminal paper presenting the potential outcomes framework, on which typical causal inference approaches are based—highlights how the results of a (true or quasi) experiment are generally of interest only to the extent that the observed data are representative of a population of future treatment assignments, i.e., that the causal relationship has a certain predictive ability. This is especially relevant in the case of climate impacts. The answer to "*Do extreme climate events lead to conflict or migration?*" has become an increasing concern under the assumption that it tells us something about "*Will future climate changes bring more conflict or migration?*".

This underlying motivation is revealed by many causal inference studies themselves, which include predictive statements or pair the estimates from explanatory causal models with climate model output to form projections. In the latter case, Carleton et al. (2016)¹³ aptly describe the need to account for the multiple sources of uncertainty, namely (i) the statistical uncertainty from

the fitted model, (ii) the variation in climate model predictions, and (iii) the potential adaptation of societies to climate change that could alter the response function. We emphasize the concern that the original fitted model might have little predictive power if it uses only climate regressors and fixed effects while imposing strong assumptions on unmodeled group-level effects.

In analyses concerned with estimating causally-interpretable parameters, considering the predictive ability of the model would not only support the external validity of the analysis as well as an ultimate motivation of the research, but also bolster its internal validity. This has been notably illustrated in the study of conflict²⁴. Prediction accuracy supports causal inference by providing an additional check on its assumptions, namely: the statistical assumptions about the data-generating process, and the aforementioned assumption of *subjective random sampling* of trials. Modeling assumptions about the data-generating process can be supported in the pre-modeling phase by using prior theory to dictate the model, but also post-modeling, by testing the model against reality, i.e., by assessing its predictive accuracy²⁵. As social science knowledge can often be too limited for deriving precise specifications, prediction provides a way to evaluate whether these assumptions hold in practice.

In summary, in the climate-migration-conflict nexus, the features of the data-generating process and the importance of predictive interpretations motivate both a systematic examination of the fixed-effects linear model assumptions and the exploration of more flexible approaches. In the next sections, we explore leveraging the same quasi-experimental variation in climate through a framework that addresses these complementary statistical goals. We first introduce the implementation of the multilevel Bayesian framework with such causal identification strategies. We describe how it accommodates various conditional distributions of outcomes, models dependence in residuals and heterogeneous effects across spatial units, and propagates uncertainty into projections under simulated climate conditions.

The multilevel Bayesian framework

This section introduces the multilevel Bayesian framework as a generalization of the fixed-effects model. We first restate the fixed-effects specification, recasting it within this framework to highlight its underlying assumptions, and then show how a multilevel structure with partial pooling relaxes them.

Longitudinal data have a hierarchical structure, where repeated measures at the lower level (time within group) are nested within higher-level group data, for example, regions. Recast in a hierarchical (i.e., multilevel) modeling framework, the MLR-FE model (1) fit to such data embeds restrictive assumptions about between-group variation: regional intercepts are completely independent or *unpooled*, while climate variables have homogeneous or *fully-pooled* effects across regions. This *varying-intercepts fixed-slopes* model is represented in equation (2), where the conditional distribution of the outcome variable is generalized as $\mathcal{F}(\mu, \theta)$ with mean μ and other parameters θ , and we consider the example of a vector of two climate-related predictors, such that $\beta = (\alpha, \gamma)'$ and $W_{it} = (A_{it}, B_{it})'$:

$$y_{it} \sim \mathcal{F}(\mu_{it}, \theta), \quad \mu_{it} = \alpha A_{it} + \gamma B_{it} + X'_{it} \delta + \phi_i + \psi_t; \quad i = 1, \dots, n; \quad t = 1, \dots, T. \quad (2)$$

In this fixed-effects model, each regional intercept ϕ_i is estimated from the data of the given region i only, which is equivalent to assuming that the intercepts belong to a joint distribution with an infinite variance. Homogeneous slope parameters, on the other hand, implicitly represent the other extreme of zero variance between regions. It is often reasonable to assume instead some degree of closeness between effects across regions, and hence use a compromise between *full pooling* and *no pooling* of regional coefficients. In effect, one may assume that they belong to a common distribution and let the degree of pooling be determined by the data. This is easily implemented with a multilevel model, where within-group variation is explicitly modeled at the lower level, and between-group variation at the higher level. Modifying model (2) to allow for such *partial pooling* of both slopes and intercepts across regions results in the two-level model (3), where the vector of region-level coefficients is assumed to follow a joint multivariate normal (MVN) distribution. Group averages of the causal variables are added as covariates to address concerns of bias from time-invariant confounders introduced by the modeled region-level effects. This model is sometimes referred to as the "within-between random effects" model²⁶ or the "correlated random effects" model²⁷. By including the group averages as regressors, the problematic correlation between the treatment variables and the group effects is removed from the group-level error term, resolving the concern of bias from group-level (here: time-invariant) confounders in a manner similar to the fixed-effects model²⁸.

$$\begin{aligned} y_{it} &\sim \mathcal{F}(\mu_{it}, \theta), \quad \mu_{it} = a_i A_{it} + b_i B_{it} + X'_{it} \delta + f_i + \eta_1 \bar{A}_i + \eta_2 \bar{B}_i + \psi_t; \quad i = 1, \dots, n; \quad t = 1, \dots, T \\ \begin{bmatrix} a_i \\ b_i \\ f_i \end{bmatrix} &\sim \text{MVN} \left(\begin{bmatrix} \alpha_0 \\ \gamma_0 \\ \phi_0 \end{bmatrix}, \begin{bmatrix} \sigma_a^2 & \sigma_{ab} & \sigma_{af} \\ \sigma_{ab} & \sigma_b^2 & \sigma_{bf} \\ \sigma_{af} & \sigma_{bf} & \sigma_f^2 \end{bmatrix} \right) \end{aligned} \quad (3)$$

In addition to relaxing the constraints on the distribution of regional coefficients, the multilevel structure allows them also to be informed by group-level predictors. In this simplest form of the model, the serial or spatial dependence structure is not

explicitly represented; however, it could be modeled by including relevant additional regressors—such as lags of the dependent variable—alongside X_{it} .

This model can be estimated in a frequentist or Bayesian framework. We emphasize the Bayesian framework because it provides full posterior distributions for group-level and population parameters, regularizes estimation in small samples, and facilitates evaluation of predictive performance through posterior predictive checks. Specifically in the climate-conflict-migration context, it simultaneously allows for: (i) accounting for the typically limited nature of migration and conflict outcome data through an adequate choice of $\mathcal{F}()$, (ii) partially pooling the intercept and slope coefficients across groups efficiently, and (iii) propagating uncertainty in parameter estimation. The posterior distributions produced can indeed be combined with simulations from climate models to account for the uncertainty in estimation when projecting outcomes.

The same identification strategies based on quasi-experiments and adjustments for confounders can be leveraged within this framework, which generalizes the typical linear model. Like any statistical approach, the multilevel Bayesian model also rests on assumptions: the functional form of the model, the specification of the outcome distribution, and the priors assigned to parameters all shape the resulting inferences. It is therefore essential to question these assumptions, and use tools such as posterior predictive checks to assess their plausibility. This framework also implies tradeoffs. The most prominent are higher computational costs and a potential reduction in the effective degrees of freedom for statistical tests. A further, though presumably less consequential, tradeoff is that of one linear assumption for another. Indeed, with a *non-identity* link function, the bias from time-invariant confounders is fully removed by including the group average of the causal variable as a covariate only if the random effect is a linear function of this average. That being said, simulations suggest that the remaining bias remains small in most situations²⁶, and additional functions of the group average can also be included to characterize more flexible functional forms of the correlation. Although these tradeoffs exist, we propose that when model diagnostics show strong departures from the assumptions of the typical linear model—such that t tests of the coefficients of the best linear approximation to the conditional expectation function do not provide the desired information—and when heterogeneity in treatment effects or prediction are of interest, as is generally the case in the climate-economy literature, these tradeoffs may be worthwhile.

Multilevel models have been used to study the relationship between climate and migration. For example, Nawrotzki et al. (2015)²⁹ use a two-level regression model to account for the hierarchical structure of their data (households nested in municipalities). In a similar setting, Nawrotzki et al. (2013)³⁰ consider a third level and include state-level predictors, explicitly addressing the fact that “migration decisions are influenced by forces operating at different scales.” In both studies, the central findings are the variation of the climate-migration association by location characteristics. However, only intercepts are modeled as random effects; slopes are fully pooled across units. The estimated relationships are not causally interpretable because they lack a strong identification strategy, and the predictive ability of the model is also not assessed. In the context of conflict outcomes, Burke et al. (2015)¹² use a hierarchical Bayesian model to conduct a meta-analysis of estimates from the literature, themselves selected for their use of the MLR-FE framework. In the present paper, we suggest instead using the multilevel Bayesian framework in combination with established identification strategies, to combine the focus on identification with the consideration of modeling assumptions and account for heterogeneity by modeling slope parameters in addition to intercepts. This framework can be adopted not solely in settings with hierarchical levels within the spatial dimension, but in any longitudinal data.

In the next section, we apply this model to a dataset representative of the literature and compare the insights obtained with those from the fixed-effects linear model.

Application: Temperature anomalies and civil war in Somalia

We consider the dataset on climate and conflict in Somalia used in Maystadt and Ecker (2014)³¹, hereafter M&E. This choice is motivated by the availability and representativeness of these data: their analysis was conducted within the MLR-FE modeling framework, and has been cited in a substantial number of literature reviews published since its original publication^{3,4,6,12–14,32}.

The hypothesis originally tested with this dataset is that temperature extremes are an indirect determinant of conflicts in Somalia operating primarily through the channel of livestock prices. The dataset contains longitudinal, monthly data at the scale of administrative regions, over the period 1997–2009. Supplementary Figure S4 online shows a map of these 18 regions. The causal identification strategy uses the exogeneity of two main explanatory variables capturing the level and duration of temperature extremes—‘temperature anomaly’ TA and ‘drought length’ DL —and controls for ‘precipitation anomaly’ PA . The outcome of interest is the number of violent conflict events. The main variables are listed in Table 1 along with the raw data sources and transformation steps. The dataset also contains livestock prices, instrumented by the climate variables in a two-stage least-squares fixed-effect model to explore the mechanism driving the reduced-form relationship. As that instrumental variable specification adopts the same functional form and treatment of the error terms as the reduced-form—and our study focuses on these modeling assumptions, not causal-identification assumptions—we present here only the reduced-form model.

Inferences from the MLR-FE model and other single-level specifications

We first estimate the effects of the climate variables of interest using the MLR-FE model, presented in equation (4), where i refers to the region, m the calendar month, y the year, and e_{imy} represents the error term.

$$\begin{aligned} conflict_{imy} &= \alpha TA_{imy} + \gamma DL_{imy} + \delta PA_{imy} + \phi_i + \psi_{my} + \omega_{im} + e_{imy}, \quad e_{imy} \sim \text{normal}(0, \sigma); \\ i &= 1, \dots, 18; \quad m = 1, \dots, 12; \quad y = 1, \dots, 13 \end{aligned} \quad (4)$$

Like M&E, we include 18 region fixed effects captured by ϕ_i , 156 month-year fixed effects ψ_{my} , and 216 region-month fixed effects ω_{im} . Temporal and spatial dependencies are accounted for simultaneously by estimating the variance-covariance matrix of the error term following the methods of Newey and West (1987)²⁰ and Conley (1999)¹⁹ with uniform weighting kernels. Spatial dependency is assumed to disappear beyond a cutoff point of 263 kilometers, corresponding to the maximum distance between the centroids of any pair of neighboring regions, and time dependency is allowed up to four months. The summary of the results is presented in the first two columns of Table 2, and diagnostic plots of the residuals are presented in the first row of Figure 1. Additional diagnostic plots are provided in detail in the Supplementary Information S.2 online.

The normal quantile-quantile diagnostic plot shows a strong departure from normality in the distribution of the residuals. A Shapiro-Wilk test and a Kolmogorov-Smirnov test further confirm this departure. TA and DL have a correlation coefficient of 0.28 when fixed effects are removed (see Supplementary Fig. S1 online). The computation of the correlations between explanatory variables by region shows that PA and DL are highly correlated in the northern regions of Awdal, Nugaal, Sool, and Togdheer. The share of the variation in outcomes within regions that is explained by the model, captured by the *within R*², is negligible when adjusted for the degrees of freedom. The correlations of climate variables, the non-normality of the residuals, and the lack of information on the true correlation structure of the residuals, render the *p*-values of the regression coefficients unreliable indicators for causal identification.

We therefore explore the robustness of the results to an alternative model specification better suited to the data-generating process. We consider a negative binomial (NB) conditional distribution to account for the count nature of the outcome data, its positive skew and the frequency of zero values:

$$\begin{aligned} conflict_{imy} &\sim \text{NB}(\mu_{imy}, \Theta); \quad i = 1, \dots, 18; \quad m = 1, \dots, 12; \quad y = 1, \dots, 13 \\ \mu_{imy} &= \alpha TA_{imy} + \gamma DL_{imy} + \delta PA_{imy} + \phi_i + \psi_{my} \end{aligned} \quad (5)$$

The second row of Figure 1 presents the diagnostic plots of this NB model (5), estimated as a generalized linear model with logarithmic link function. As expected, we no longer observe a stark departure from the modeling assumptions; the Akaike information criterion (AIC) is also substantially lower. With this distribution, we find that the key climate variables of interest are not statistically significant (Table 2).

A multilevel Bayesian model of climate-fueled conflict

We now relax the assumptions of no-pooling of baseline effects and full pooling of climate effects across regions by generalizing to the multilevel negative binomial model (6), which partially pools slopes and intercepts across regions, and we estimate it in the Bayesian framework. Removing the region-month interaction term ω_{im} present in the single-level model specifications gives comparable results and is more computationally efficient, so we omit it from the multilevel model for simplicity. All parameters, including the hyperparameters of the between-group covariance matrix, are assigned *weakly informative* priors, i.e., prior distributions that mildly constrain parameters toward plausible values to regularize estimation without overwhelming the information in the data (see the Supplementary Information S.4 online for details).

$$\begin{aligned} conflict_{imy} &\sim \text{NB}(\mu_{imy}, \Theta); \quad i = 1, \dots, 18; \quad m = 1, \dots, 12; \quad y = 1, \dots, 13 \\ \mu_{imy} &= a_i TA_{imy} + b_i DL_{imy} + c_i PA_{imy} + f_i + \eta_1 \overline{TA}_i + \eta_2 \overline{DL}_i + \eta_3 \overline{PA}_i + \psi_{my} \\ \begin{bmatrix} a_i \\ b_i \\ c_i \\ f_i \end{bmatrix} &\sim \text{MVN} \left(\begin{bmatrix} \alpha_0 \\ \gamma_0 \\ \delta_0 \\ \phi_0 \end{bmatrix}, \begin{bmatrix} \sigma_a^2 & \sigma_{ab} & \sigma_{ac} & \sigma_{af} \\ \sigma_{ab} & \sigma_b^2 & \sigma_{bc} & \sigma_{bf} \\ \sigma_{ac} & \sigma_{bc} & \sigma_c^2 & \sigma_{cf} \\ \sigma_{af} & \sigma_{bf} & \sigma_{cf} & \sigma_f^2 \end{bmatrix} \right) \end{aligned} \quad (6)$$

Figure 2 shows the posterior distributions of the central slope coefficients of the two negative binomial models— (α, γ, δ) for the single-level model (5) and $(\alpha_0, \gamma_0, \delta_0)$ for the multilevel model (6)—estimated using a Bayesian framework. For each parameter of interest, Bayesian estimation produces a full posterior distribution, which represents the updated uncertainty about the value of the parameter after observing the data. To compare this estimation to results from a frequentist approach, like model (5) (which produces a single point estimate with an associated standard error and confidence intervals), we compute

summary statistics of the posterior distribution, specifically the median as a measure of central tendency and 95% credible intervals. As expected, the median values of the distributions from the fixed-slope Bayesian model are virtually identical to the frequentist point estimates. However, in the partial pooling model, the distribution of the mean slope coefficient for DL shifts towards zero, making the evidence of an effect of the climate variable is inconclusive.

The posterior distributions of the partially-pooled slope coefficients (a_i, b_i, c_i) provide some insight into what drives these higher-order effects. Figure 3 shows substantial heterogeneity in the effect of DL between regions, with only a few southern regions (Banaadir, Bay, and Gedo) actually driving the positive causal relationship, while others experience negative effects, which helps explain the relatively weak results in the full-pooling model. The Banaadir and Bay regions experienced the lowest numbers of months without any conflict throughout the study period. The uncertainty around the parameters also shows large differences between regions. In the northern regions of Awdal and Sool, the large uncertainty reflects the low total number of conflicts over the study period (one and four, respectively).

Predictive ability. Considering the model's predictive ability—whether for model evaluation or for simulation—is hindered by the presence of temporal fixed effects. Indeed, what are the values of the coefficients on year-month indicator variables for future periods? Similarly to pooling region intercepts in place of using region fixed effects, the assumption of unrelated effects across months can be relaxed by pooling them, i.e., by modeling them as random effects. We can then form predictions for a time period outside of the sample by sampling from their estimated distribution. We explore this in model (7), where we further add predictors for the country-wide period effect w_t . We include the group averages of the causal variables (taken over the regions within each period) to address unbiasedness concerns, as well as a time-varying climate forcing C_t experienced across all regions, and model w_t as a simple linear function of this predictor. Given a climate time series c_{it} , such as a drought index—or in our case the precipitation anomaly PA_{it} —we construct the climate forcing predictor $C_t \equiv \sum_i \frac{c_{it}}{\mathbb{Q}_{c_i}(.5)}$ which emphasizes threshold exceedance. Here $\mathbb{Q}_{c_i}(.5)$ is the median of the climate time series. The general motivation for including group-level predictors is that they may reduce unexplained group-level variation and thus yield more precise estimates than by shrinking all groups equally toward the population mean. For this specific predictor, the motivation is two-fold: we may want to assess whether conflict responds to a pervasive spatial impact or an aggregate climate anomaly aside from a regional climate anomaly, and we would expect it to be related to tail behavior of the same sign across locations. We also consider an alternative definition of the climate forcing that captures deviations from the average, with very similar results (see Supplementary Information S.3.2 online).

$$\begin{aligned}
conflict_{imy} &\sim NB(\mu_{imy}, \Theta), \quad i = 1, \dots, 18; \quad m = 1, \dots, 12; \quad y = 1, \dots, 13; \\
\mu_{imy} &= a_i TA_{imy} + b_i DL_{imy} + c_i PA_{imy} + f_i + \eta_1 \overline{TA}_i + \eta_2 \overline{DL}_i + \eta_3 \overline{PA}_i + w_{my} \\
w_{my} &\sim \text{normal}(\psi_0 + \psi_1 \overline{TA}_{my} + \psi_2 \overline{DL}_{my} + \psi_3 C_{my}, \sigma_w)
\end{aligned} \tag{7}$$

$$\begin{bmatrix} a_i \\ b_i \\ c_i \\ f_i \end{bmatrix} \sim \text{MVN} \left(\begin{bmatrix} \alpha_0 \\ \gamma_0 \\ \delta_0 \\ \phi_0 \end{bmatrix}, \begin{bmatrix} \sigma_a^2 & \sigma_{ab} & \sigma_{ac} & \sigma_{af} \\ \sigma_{ab} & \sigma_b^2 & \sigma_{bc} & \sigma_{bf} \\ \sigma_{ac} & \sigma_{bc} & \sigma_c^2 & \sigma_{cf} \\ \sigma_{af} & \sigma_{bf} & \sigma_{cf} & \sigma_f^2 \end{bmatrix} \right)$$

Figure 4 shows a graphical comparison of relevant statistics of the observed data—namely, the proportion of zeros, due to the count nature of the outcome—against replicated datasets from the fits of the different models considered. This comparison of the posterior distributions shows that the original Gaussian model is not able to capture the count nature of the data. The negative binomial models with no pooling or partial pooling across regions provide better fits, and the partial pooling of temporal effects produces the best fit for the data. The higher performance of the multilevel models is also supported by cross-validation. We estimate the expected log predictive density (ELPD) of each model using leave-one-out cross-validation, then compute the differences between these ELPD estimates along with the standard errors of these differences. The results in Table 3 show the same hierarchy in expected predictive accuracy as in the comparison of posterior predictive statistics. We also replicate the partial-pooling model with the original Gaussian functional form, to assess how the predictive performance changes when using a multilevel model, compared to the original specification (see Supplementary Information S.3.1 online). We find, as captured by the ELPD, that the partial-pooling models perform marginally better than the no-pooling model; however, they do not match the performance of the NB model which accounts for the count nature of the outcome data.

The application of hierarchical Bayesian models to this dataset illustrates how the standard MLR-FE model can be readily recast in the hierarchical Bayesian framework, with a better suited conditional distribution for the response variable, inferences on parameters generated using simulations—which enable the propagation of uncertainty if estimates are later used in projections—and the partial pooling of regression coefficients, and can thereby allow one to analyze and model differences across groups efficiently. The posterior distributions obtained can easily be combined with simulations similar to those proposed in M&E, e.g., of increases in one standard deviation in the main climate variables, or of projected changes from global climate

models under climate scenarios, to address what we can reasonably say about questions akin to ‘*Will future climate changes bring more conflict or migration?*’ and with how much uncertainty.

Discussion

The hierarchical Bayesian approach to studying climate-fueled migration and conflict addresses several statistical principles of particular concern in this interdisciplinary field. First, prediction and identification of causal effects are complementary pursuits. Together, they support the modeling assumptions and thus the internal validity of the statistical analysis while strengthening its external validity, which is generally of practical interest. Reporting both the model’s explanatory power and its predictive power, and considering the latter for model selection, may strengthen inference from natural experiments²². Second, the linear approximation of the conditional expectation function of the outcome—which is the target of the linear reduced-form model—may be misleading when the data-generating process departs substantially from the modeling assumptions. Separately from the identifying assumptions supported by quasi-experimental variation and the ability to adjust for confounders, the validity of inferences also rests on the underlying assumptions of the estimated model. An assessment of the key assumptions such as the conditional distribution of the outcome, the absence of high influence observations, and the non-collinearity of regressors is hence of particular interest and may be provided alongside the estimation results—see for example Cohen et al. (2008)³³. Third, the response to climate variables can vary across spatial units but with the ability to estimate spatial sensitivity limited by the sample size. In such cases, a fully pooled model for the slope coefficients, as is typical in the MLR-FE framework, represents a tradeoff between the efficiency gained by using a larger sample size (by pooling across all locations) and the potential bias at each unit. Instead, generalizing to a partial pooling approach combines information across units, shrinking uncertainty while reducing bias at individual spatial units. This multilevel structure provides a principled way to assess whether variation in climate sensitivity across groups is meaningful and what observed attributes it may depend on.

The multilevel Bayesian framework, which is a generalization of the linear reduced-form model, brings the analysis closer to addressing the three concerns above, thereby strengthening inference from quasi-experiments in the climate-migration-conflict nexus. The benefits of such models and inference methods have been shown abundantly in the statistics literature and are relevant in many settings using observational data to generate causal and predictive inferences. They come down to considering more general statistical frameworks to learn from quasi-experimental data, and selecting models that approximate the data-generating process; essentially applying the recommendation of “a combination of the economists’ focus on identification strategies and the statisticians’ ability to build more complicated models to assess what might happen if the strict assumptions fall apart”³⁴. This study emphasizes and illustrates how these benefits are particularly salient for studying the relationship between climate and social instability. In this application, most of the contrast in estimated climate effects arises from adopting a negative binomial outcome distribution, which better reflects the count nature of the conflict data. The Bayesian framework complements this correction by enabling partial pooling and coherent assessment of uncertainty and predictive performance, extending the analysis to capture heterogeneity and strengthen inference within a unified structure. Our findings suggest that a multilevel Bayesian approach has the potential to support the internal validity, external validity, and efficiency of inferences, providing a robust foundation for understanding the impacts of changes in climate on social outcomes of interest. This methodological bridge between econometric identification strategies and complex statistical modeling capabilities addresses a critical need in interdisciplinary climate impact research, where collaboration across disciplinary boundaries is essential.

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Author contributions

C.P., P.C.L., and U.L. designed research and wrote the main manuscript; C.P. and P.C.L. analyzed data; all authors reviewed the manuscript.

Data and code availability

The code and datasets generated and used for the study are available on Zenodo at <https://doi.org/10.5281/zenodo.1716512>.

Additional information

Competing interests. The authors declare no competing interests.

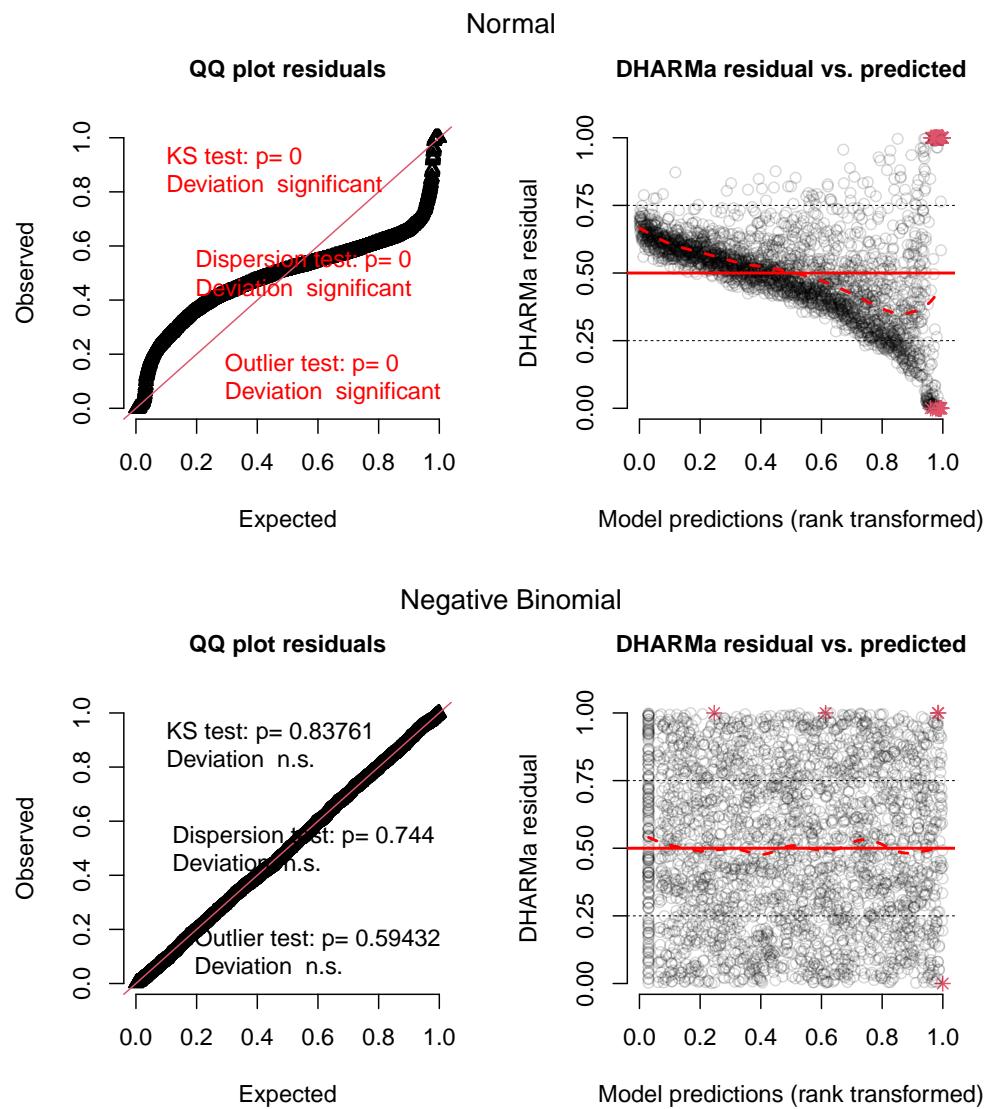


Figure 1. Residual plots of the reduced-form model assuming a conditional normal distribution (top) or negative binomial distribution (bottom). Left: quantile-quantile plot of the residuals, including Kolmogorov-Smirnov (KS) test for the goodness of fit of the residuals to the specified distribution. Right: standardized residuals against model predictions; red stars represent outliers.

Description	Name	Original resolution and processing steps	Source		
Count of violent conflict events	<i>conflict</i>	region \times month	ACLED (2011) ³⁵		
Temperature anomaly	<i>TA</i>	0.5° \times 0.5° grid \times month (average of daily maximum Ts) a. T interpolated to region centers by kriging; b. anomaly computed w.r.t. 1980–2009; c. anomaly averaged over 3 months	CRU	TS	3.1 (2008) ³⁶
Drought length	<i>DL</i>	count of consecutive months with positive <i>TA</i> values	CRU	TS	3.1 (2008) ³⁶
Precipitation anomaly	<i>PA</i>	0.5° \times 0.5° grid \times month (total precipitation P) a. P interpolated to region centers by kriging; b. anomaly computed w.r.t. 1983–2009; c. anomaly averaged over 3 months	CRU	TS	3.1 (2008) ³⁶

Table 1. Main variables in M&E’s original reduced-form specification

Conditional distribution	Normal (original ^a SEs)	Normal (corrected SEs)	NB	NB
<i>TA</i>	0.71 (0.25)	0.71 (0.37)	0.01 (0.12)	0.08 (0.11)
<i>DL</i>	0.08 (0.01)	0.08 (0.04)	0.02 (0.01)	0.02 (0.01)
<i>PA</i>	-0.47 (0.20)	-0.47 (0.31)	-0.08 (0.16)	-0.07 (0.15)
region FE ^s	✓	✓	✓	✓
year-month FE ^s	✓	✓	✓	✓
region-month FE ^s	✓	✓	✓	
<i>N</i>	2,808	2,808	2,808	2,808
multiple <i>R</i> ²		0.43		
within ^b <i>R</i> ²	0.17	0.17		
adjusted within <i>R</i> ²		0.04		
AIC		16121	5837	5639

^a The ‘original’ standard errors (SEs) are those reported in M&E. They were computed using a version of the `ols_spatial_HAC` function by Hsiang (2010)³⁷ which miscalculated the weights for serial autocorrelation. Using the corrected version (v3) in the authors’ Stata code results in the ‘corrected’ standard errors in the adjacent column.

^b The within or ‘projected’ *R*² corresponds to the *R*² of the mean-deviated regression, i.e., after removing region fixed effects, and represents the share of the variation in the outcome within regions that is captured by the model.

Table 2. Frequentist estimates of slope coefficients of the single-level models. The raw regression coefficients are displayed; they represent additive changes in *y* in Gaussian models, and additive changes in $\log(y)$ in negative binomial models, where *y* is the number of conflicts.

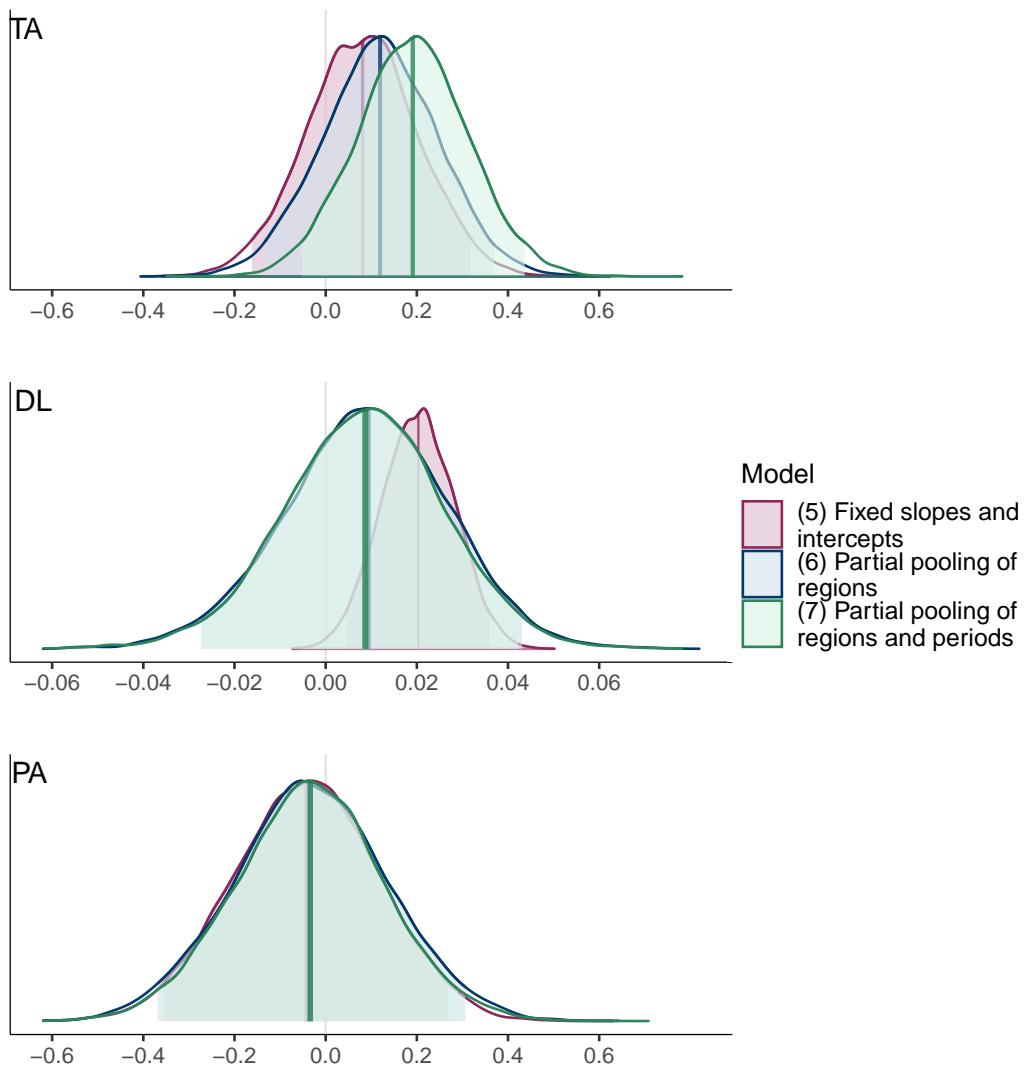


Figure 2. Posterior distributions of the central slope coefficients of the Bayesian negative binomial models. Pink: single-level model (varying-intercept, fixed-slope), parameters α, γ, δ . Blue and green: multilevel models (blue: intercepts and slopes partially pooled across regions; green: intercepts partially pooled across regions and periods, slopes partially pooled across regions), hyperparameters $\alpha_0, \gamma_0, \delta_0$. 95% credible intervals are shown as shaded areas under the curves around the median point estimates.

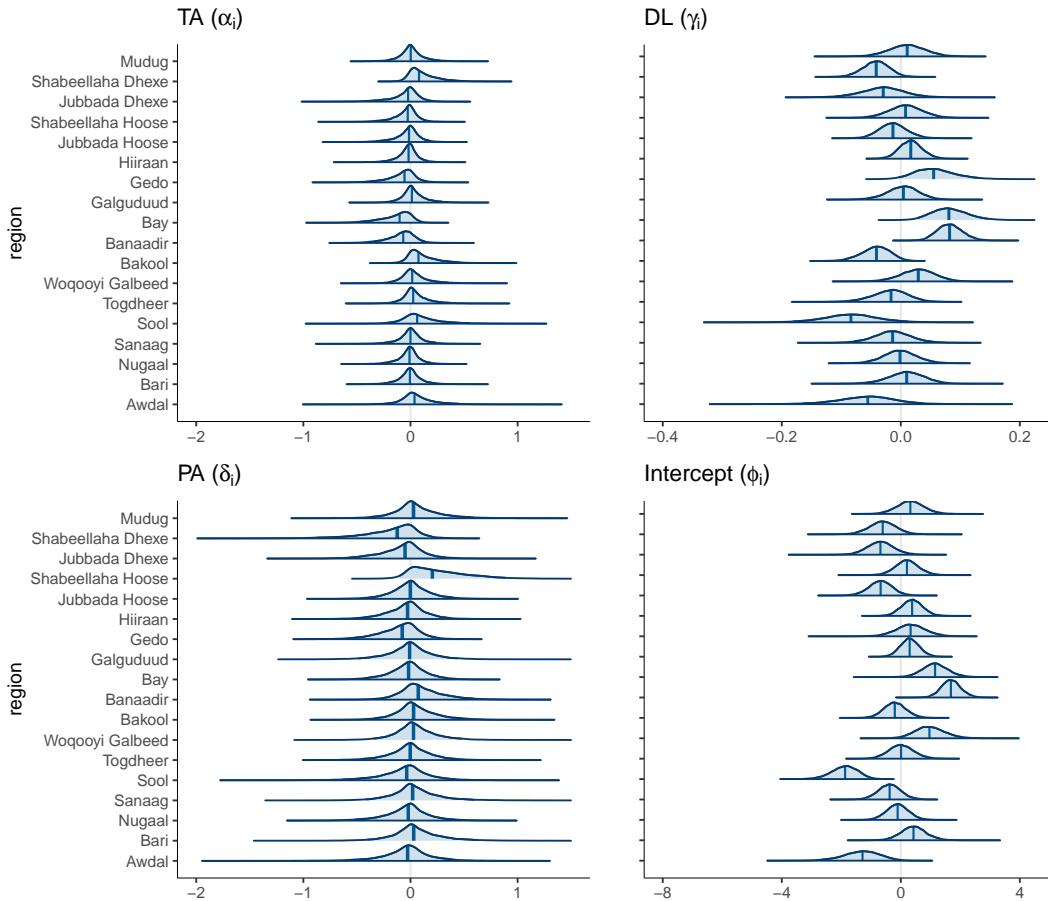


Figure 3. Posterior distributions of the partially-pooled region-specific intercepts and slope coefficients (model (6)). 95% credible intervals are shown as shaded areas under the curves around the median point estimates.

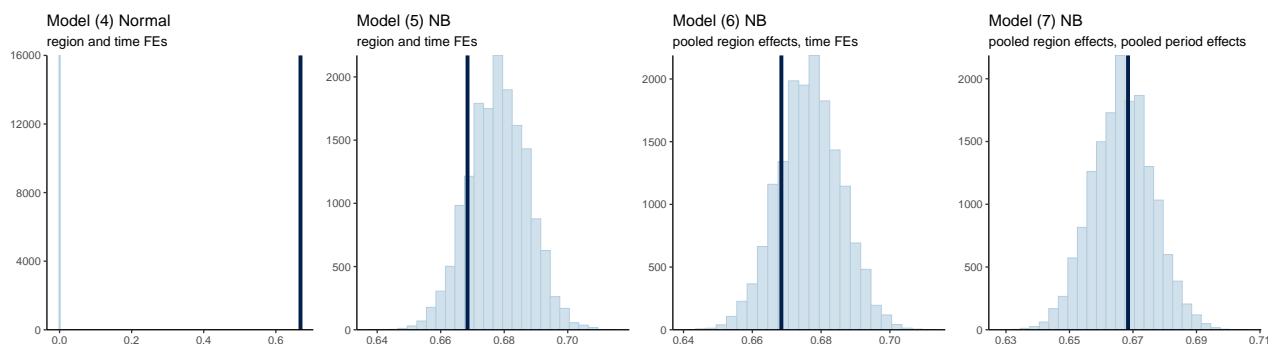


Figure 4. Graphical posterior predictive check: comparison of the proportion of zeros in observed vs simulated datasets. For each model, 8000 datasets are simulated from the posterior predictive distribution using the observed predictors. The light blue histogram represents the distribution of the value of the statistic (the proportion of zeros) across the simulated datasets. The dark blue vertical line is the value of this statistic for the observed sample.

Model	(4)	(5)	(6)	(7)
Condit. distribution	Normal	NB	NB	NB
TA	0.71 (0.30)	0.08 (0.12)	0.12 (0.13)	0.19 (0.12)
DL	0.08 (0.02)	0.02 (0.01)	0.01 (0.02)	0.01 (0.02)
PA	-0.47 (0.31)	-0.04 (0.16)	-0.03 (0.17)	-0.03 (0.16)
\overline{TA}_i			-11.6 (2.37)	-11.13 (2.23)
\overline{DL}_i			0.37 (0.19)	0.36 (0.18)
\overline{PA}_i			1.64 (3.07)	1.85 (3.07)
\overline{TA}_{my}				0.57 (0.29)
\overline{DL}_{my}				-0.07 (0.04)
\overline{C}_{my}				0.01 (0.00)
region	FEs	FEs	pooled	pooled
month	FEs	FEs	FEs	pooled
N	2,808	2,808	2,808	2,808
ELPD	-8090.2	-2861.7	-2833.8	-2809.8
ELPD _{diff}	-5280.3	-51.8	-24	0
SE[ELPD _{diff}]	186.4	13.8	9.5	0

Table 3. Bayesian central estimates of slope coefficients of single-level and multilevel models. For each regressor, the table display summaries of its marginal posterior distribution: the distribution's median (top sub-row) and an estimate of the distribution's standard deviation (bottom sub-row, in parentheses) based on a scaling of the median absolute deviation around that median. ELPD_{diff} corresponds to the difference in expected log predictive density (ELPD) between models, and SE[ELPD_{diff}] to the standard error of that difference, where the reference is the model with the largest ELPD (model (7)).

Supplementary Information

S.1 Overview of systematic reviews of the climate-migration-conflict literatures (2013-2024)

Over the past twelve years (2013-2024), the literature examining relationships between climate change, migration, and conflict has been subject to extensive systematic review, with at least 21 reviews published during this period—12 focusing on migration^{1,2,5,9,38–45} and 9 on conflict^{3,4,6,10,12–14,32,46}. This proliferation of meta-analytical work reflects both the growing importance of the topic and the persistent challenges in reaching consensus about climate impacts on human mobility and social stability.

The systematic reviews highlight a high methodological diversity. Early syntheses identified distinct modeling approaches including spatial vulnerability models, hazard analysis models, and agent-based modeling for migration studies^{40,43}. The econometric literature has favored multivariate regression models and gravity models, while simulation approaches have employed agent-based models and integrated assessment models. The inference approaches used also differ. At the highest level, we can first separate two classes of models: (i) computational models, which use simulations to study systems of interest; these include agent-based models^{47,48}, system dynamics models⁴⁹, computable general equilibrium models⁵⁰, and integrated models⁵¹; (ii) empirical analyses using regression models, from simple least squares to multilevel generalized linear models and nonparametric predictive models. Regression-based approaches can be classified along multiple dimensions, such as the functional form of the relationship between predictor and outcome variables, the assumed conditional distribution of the outcome variable, and the stated end goal of the statistical exercise—notably prediction or causal inference (which can be considered as conditional prediction). Table S1 illustrates the diversity of approaches as they leverage some combination across these key dimensions, based on a sample of studies in the literature—each cited by at least one of the 21 literature reviews. Multiple reviews also emphasize that the lack of comparable data and standardized measures across contexts prevents making generalizations that would facilitate the development of future scenarios^{5,40}. The literature does not use core concepts and variables in a standardized manner, with different operationalizations of dependent variables (onset versus incidence of conflict, various types of displacement) and independent variables requiring nuanced theoretical approaches³.

The migration literature presents complex findings. Reviews consistently identify that environmental change alone does not cause migration; rather, it is the interaction of environmental factors with non-environmental factors that drives mobility decisions^{5,39}. This insight has led to the development of frameworks distinguishing between vulnerability and capability effects, where climate shocks affect both household resources necessary for migration and the risk of remaining in place⁴¹. Several important patterns emerge: climate variability produces different responses than long-term climate change, slow-onset events produce different migration outcomes than fast-onset events, and the relationship between climate severity and migration is non-linear, with threshold effects determining whether vulnerability or capability channels dominate. The most recent and largest meta-analysis synthesizes 96 studies and finds an overall average effect on migration that is small for both slow- and rapid-onset events, but positive and significant⁹. It also highlights that across studies, climate is found to play diverse and even opposite roles: a driver of migration, a constraint to mobility, or having no significant role among other drivers—highlighting the fundamental heterogeneity in climate-migration relationships.

The conflict literature presents a more polarized landscape. Early systematic reviews found little consensus on the physical pathways connecting climate change to violent conflicts^{3,52}. However, the emergence of the "climate-economy" literature, employing reduced-form econometric models with high-frequency climate variation and fixed effects, has claimed to establish robust causal relationships. Meta-analyses that examined studies meeting strict econometric criteria—requiring time-series variation, location-specific fixed effects, and trend controls—found consistent support for climate-conflict links^{12,14}. However, reviews also note a frequency-identification trade-off: populations evolve faster than many low-frequency climatic changes of interest, undermining the comparability of control and treatment populations over time^{12,13}, and highlight the limited insights into mechanisms and policy pathways⁶. The latest review still notes "a remarkable inconsistency of evidence among publications"¹⁰.

Across the climate-migration-conflict nexus, a critical finding across assessments is the presence of significant methodological limitations and selection biases. Methodological choices substantially influence the probability of finding climate effects, with different econometric techniques, climate measures, and outcome definitions producing systematically different results¹. The literature's focus on econometric identification has excluded valuable predictive modeling approaches^{1,41}, while geographic coverage shows systematic biases with certain regions over-represented while other conflict-affected or climate-vulnerable regions remain understudied. The reviews consistently identify the absence of robust theoretical frameworks as a fundamental limitation. This theoretical gap contributes to the proliferation of empirical approaches without clear guidance on appropriate model specifications or expected effect magnitudes. A recurring concern is also the external validity of historical relationships for future climate scenarios, as adaptation could fundamentally alter observed relationships^{13,32}.

Conclusion. After more than a decade of systematic review, the persistence of mixed findings across multiple comprehensive reviews suggests that methodological diversity, while valuable, may be insufficient to resolve core empirical questions in the

climate-migration-conflict literature without stronger theoretical foundations and more attention to predictive performance alongside causal identification. This suggests that purely econometric approaches to causal identification may need to be complemented by more flexible modeling frameworks that can accommodate heterogeneity, non-linearity, and complex interactions while maintaining rigorous standards for causal inference. This integration of econometric identification strategies with predictive modeling—as advocated in our main analysis—represents a promising direction for advancing both scientific understanding and policy relevance in this critical area of research.

S.2 Diagnostic plots of the fixed effects linear model

Figure S1 presents scatterplots of the model's main variables—adjusted for the region, month-year and region-month fixed effects by using the residuals of a linear regression of each variable on these fixed effects. The results suggest no obvious strong relationships between the outcome and the explanatory variables.

Figure S2 presents diagnostic residual plots of the model. From left to right, top to bottom, the sub-panels provide the following information:

- "Residuals vs. Fitted" shows whether the residuals are equally spread around a horizontal line or display a pattern—suggesting a nonlinear relationship that was not captured by the model.
- "Normal Q-Q" shows how the distribution of the residuals aligns with a normal distribution.
- "Scale-Location" presents the spread of transformed residuals across the range of predicted values, a uniform vertical spread indicating uniform variance.
- "Residuals vs. Leverage" highlights influential observations as those outside "Cook's distance lines," i.e., whose Cook's distance (the change in the predicted value if the given observation were omitted) is large.

The heteroscedasticity in residuals raises the concern that a subset of influential observations may have a disproportionate impact on the coefficients. The *Residuals vs. Leverage* plot shows residuals with high values (up to 15 standard deviations) but not high leverage.

S.3 Alternative multilevel specifications

S.3.1 Multilevel Gaussian model

To assess how the predictive performance changes when using a multilevel model, compared to the original specification, we fit, considering the original Gaussian functional form, the different multilevel models considered, namely: pooling regional intercepts only (model (4')) and pooling separately both regional intercepts and time periods (model (4'')). Results are presented in Table S2. We find, as captured by the expected log predictive density (ELPD), that the partial pooling models perform marginally better than the model with fixed effects. However, they do not match the performance of the NB model which accounts for the count nature of the outcome data.

S.3.2 Pooling of period effects

Model (7) relaxes the assumption of unrelated effects across months by pooling the temporal effects, i.e., modeling them as random effects, and includes a time-varying linear climate predictor C_t that is experienced across all regions. This predictor for the country-wide period effect w_t is constructed from a relevant climate time series c_{it} , such as a drought index, or in our case, as it is readily available, the precipitation anomaly PA_{it} .

In our main specification, the climate predictor is defined as $C_t \equiv \sum_i \frac{c_{it}}{\mathbb{Q}_{c_i}(.5)}$, where $\mathbb{Q}_{c_i}(.5)$ is the median of the climate time series, so as to capture threshold exceedance. Alternatively, it could be defined so as to emphasize deviations from the average, such as $C_t^{\text{alt}} \equiv \sum_i \frac{c_{it} - \mu_{c_i}}{SD[c_i]}$, where μ_{c_i} is the mean of the climate time series and $SD[c_i]$ its standard deviation.

Figure S3 shows a graphical comparison of relevant statistics of the data observed—namely, the proportion of zeros, owing to the count nature of the outcome—against that of replicated datasets from the two model fits. The result suggests that the two models have comparable expected predictive ability. The estimation of their ELPD by leave-one-out cross-validation confirms this result, with a small and non-statistically significant difference of 1.2 in expected ELPD.

S.4 Estimation details

All estimations were conducted using the statistical software R (v4.5.1).

Single-level models. The linear regression for the replication exercise was estimated using the function `felm` of the `lfe` package (v3.1.1). The adjustment for serial and spatial autocorrelation, implemented in the original paper with the `ols_spatial_HAC` function available for Stata by Hsiang (2010)⁵³, was implemented by adapting the function `ConleySEs` available for R by Darin Christensen^{54,55}.

Residual plots were obtained with the R package `DHARMa` (v 0.4.7) which uses a simulation-based approach to create interpretable scaled residuals from generalized linear mixed models⁵⁶.

Bayesian models. All Bayesian models were estimated using the `rstanarm` package (v2.32.1), which interfaces with the Stan C++ library for Bayesian estimation. Each model was run with 4 chains, with 8000 iterations per chain (4000 warm-up). Convergence was assessed and ensured based on the potential scale reduction factor on split chains R-hat.

Each parameter was assigned the package's default "weakly informative" priors: centered autoscaled normal for slope coefficients and exponential(1) for the negative binomial dispersion parameter. These priors regularize estimation by ruling out extreme parameter values while leaving most likelihood information intact. Regression coefficients are given independent normal(0, 2.5) baseline priors, whose variances are then automatically adjusted to the scale of each predictor and outcome. Estimation is thereby primarily driven by the data and the partial pooling structure. The effective priors of the regressors in models 5 to 7 are presented in Table S3. For further details on the defaults, we refer the reader to the `rstanarm` documentation⁵⁷.

Because no external prior information was available, we chose weakly informative priors to minimize their influence. In such cases, the specific distributional form (e.g., normal or uniform) has little effect once the variance is set sufficiently large. More informative priors could be useful, especially with smaller or noisier datasets, if prior knowledge exists—for example, if posterior distributions from historical analyses were available. The Bayesian framework readily accommodates this flexibility, though such extensions are beyond the scope of the present study.

As estimation proceeds by simulation rather than analytic inversion of large matrices, numerical stability is not a concern. Convergence issues may arise if overly complex correlation structures were imposed in the priors.

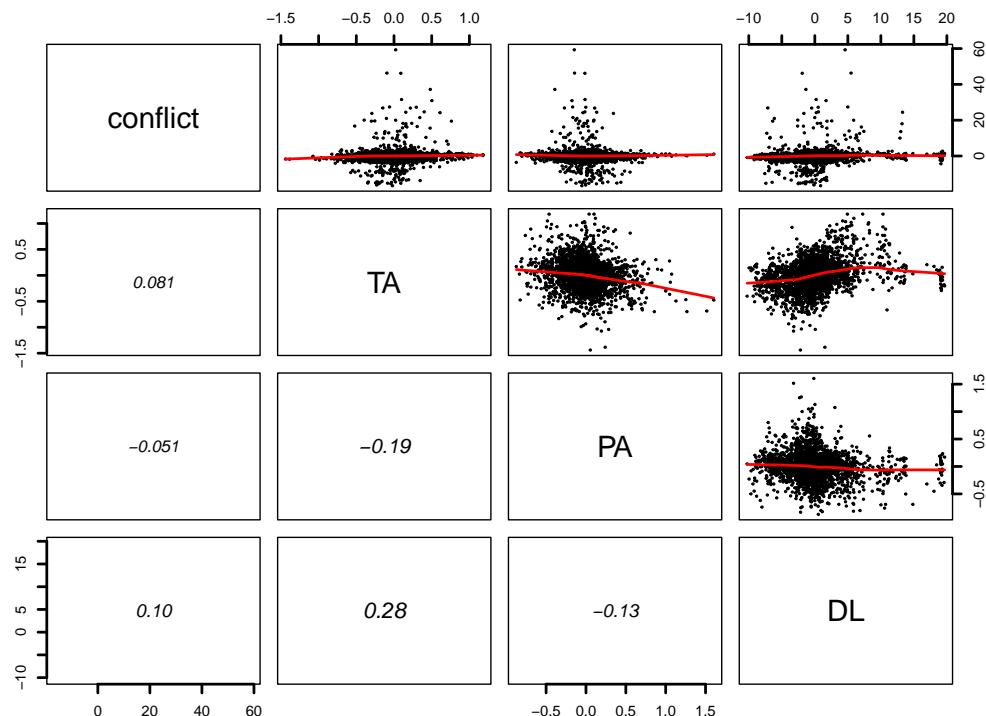


Figure S1. Scatterplots and correlations of the main variables after accounting for the region, month-year and region-month fixed effects. Top-panel: locally weighted scatterplot smoothing (LOWESS) lines with span=2/3.

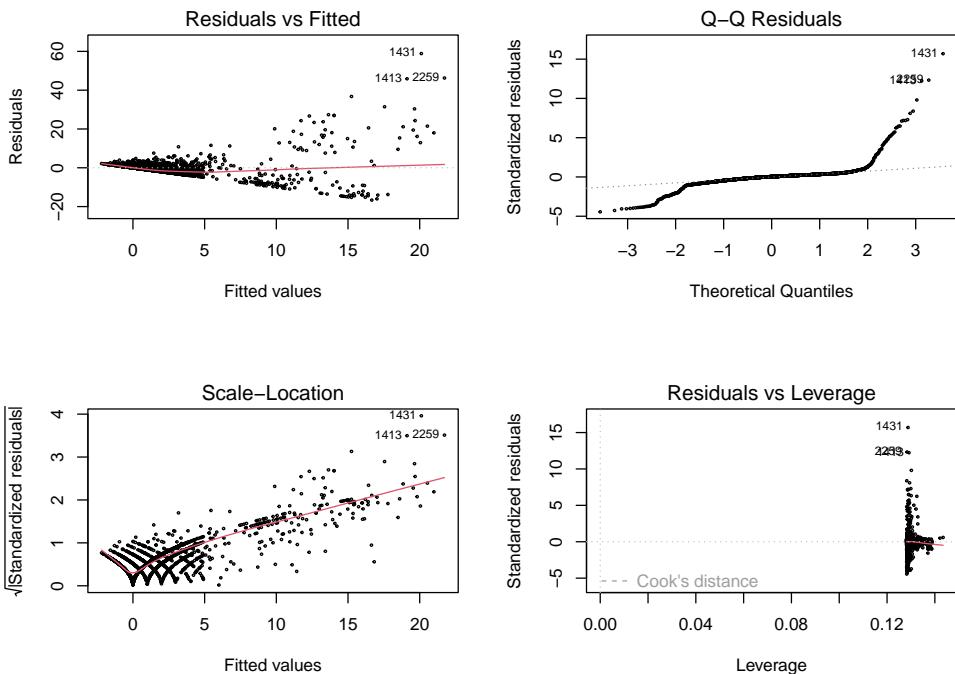
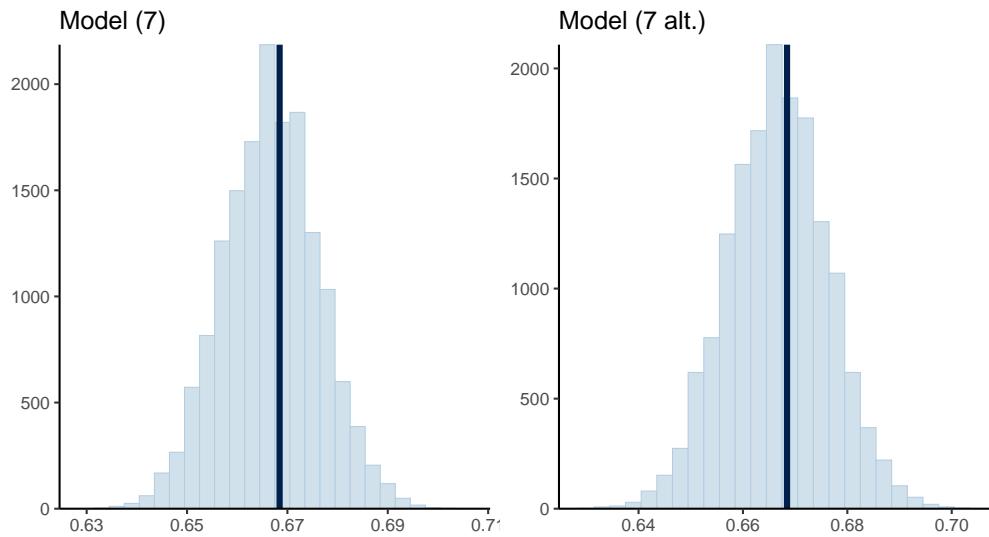


Figure S2. Diagnostic residual plots of the reduced-form model.



Notes: For each model, $S = 8000$ datasets are simulated from the posterior predictive distribution using the observed predictors. The light blue histogram represents the distribution of the value of the test statistic (the proportion of zeros) across the simulated datasets. The dark blue vertical line is the value of this statistic for the observed sample. The climate forcing \bar{C}_{my} is defined as $\sum_i \frac{c_{imy}}{\bar{Q}_{ci}(.5)}$ in model (7) and $\sum_i \frac{c_{imy} - \bar{C}_{ci}}{SD[c_i]}$ in model (7 alt.).

Figure S3. Graphical posterior predictive check: comparison of the proportion of zeros in observed vs simulated datasets; Models with partially-pooled period effects.

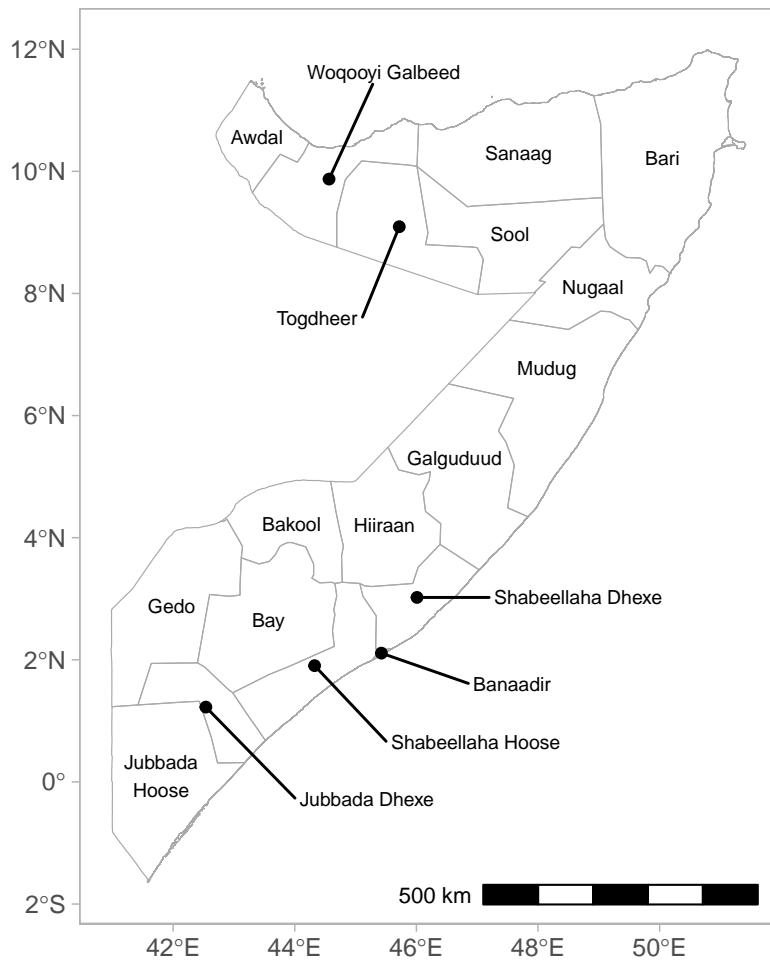


Figure S4. Map of the 18 administrative regions in Somalia. This figure was created using the statistical software R, version 4.5.1 (<https://www.r-project.org>).

	Backhaus et al. (2015) ¹⁸	Berlemann and Tran (2020) ⁵⁸	Caruso et al. (2016) ⁵⁹	Hendrix and Salehyan (2012) ⁶⁰	Henry et al. (2004) ⁶¹	Maystadt and Ecker (2014) ³¹	Nawrotzki et al. (2013) ³⁰	Nawrotzki et al. (2015) ²⁹
Pooling of individual effects								
fully pooled (no FEs)		x	x	x				
no pooling (unit/group FEs)	x	x		x		x		
partial pooling (multilevel)						x	x	
Partial vs total effect of climate								
reduced form w/o channels		x	x	x	x	x		
disentangle the reduced form relationship	x	x	x			x	x	x
Functional form of the relationship with climate explanatory variables								
linear	x	x	x	x	x	x	x	x
non-linear				x			x	x
Type of outcome variable								
binary (event)				x	x		x	x
count or flow	x	x	x	x		x		
Goal and use								
association	x				x		x	x
prediction						x		
causal inference		x	x	x		x		
Estimation								
OLS, GLS	x	x	x		x	x		
ML		x		x			x	x
GMM						x		
Bayesian					x			
Post-estimation standard error adjustment								
Heteroskedasticity robust	x				x			
Clustering			x	x				
Serial/spatial correlation using a kernel						x		

Table S1. Attributes of regression-based approaches in the climate-fueled migration and conflict literatures, based on a sample of studies selected in systematic reviews.

Model	(4)	(4')	(4'')	(5)	(6)	(7)
Condit. distribution	Normal	Normal	Normal	NB	NB	NB
TA	0.71 (0.30)	0.94 (0.31)	0.99 (0.31)	0.08 (0.12)	0.12 (0.13)	0.19 (0.12)
DL	0.08 (0.02)	0.07 (0.05)	0.07 (0.04)	0.02 (0.01)	0.01 (0.02)	0.01 (0.02)
PA	-0.47 (0.31)	-0.26 (0.29)	-0.21 (0.29)	-0.04 (0.16)	-0.03 (0.17)	-0.03 (0.16)
\overline{TA}_i		-9.33 (2.53)	-7.73 (2.57)		-11.6 (2.37)	-11.13 (2.23)
\overline{DL}_i		0.29 (0.21)	0.22 (0.21)		0.37 (0.19)	0.36 (0.18)
\overline{PA}_i		-0.59 (3.40)	0.01 (3.41)		1.64 (3.07)	1.85 (3.07)
\overline{TA}_{my}			-0.50 (0.39)			0.57 (0.29)
\overline{DL}_{my}			-0.07 (0.04)			-0.07 (0.04)
\bar{C}_{my}			0.01 (0.01)			0.01 (0.00)
region	FEs	pooled	pooled	FEs	pooled	pooled
month	FEs	FEs	pooled	FEs	FEs	pooled
N	2,808	2,808	2,808	2,808	2,808	2,808
ELPD	-8090.2	-7815.7	-7787.3	-2861.7	-2833.8	-2809.8
ELPD _{diff}	-5280.3	-5005.9	-4977.5	-51.8	-24	0
SE[ELPD _{diff}]	186.4	204.5	213.3	13.8	9.5	0

Table S2. Bayesian central estimates of slope coefficients of single-level and multilevel models. For each regressor, the two sub-rows display summaries of its marginal posterior distribution: the distribution's median (top) and an estimate of the distribution's standard deviation (bottom, in parentheses), based on a scaling of the Median Absolute Deviation around that median. ELPD_{diff} corresponds to the difference in ELPD between models, and SE[ELPD_{diff}] to the standard error of that difference, where the reference is the model with the largest ELPD (model (7)).

Model(s)	Variable(s)	Regressor(s)	Distribution
(5), (6), (7)	α_0	TA	normal(0,4.21)
(5), (6), (7)	γ_0	DL	normal(0,0.43)
(5), (6), (7)	δ_0	PA	normal(0,5.76)
(5)	$\{\phi_i\}$	$\{\mathbb{1}[region = i]\}$	normal(0,10.91)
(5), (6)	$\{\psi_{my}\}$	$\{\mathbb{1}[year, month = my]\}$	normal(0,31.32)
(6), (7)	η_1	\overline{TA}_i	normal(0,24.48)
(6), (7)	η_2	\overline{DL}_i	normal(0,1.99)
(6), (7)	η_3	\overline{PA}_i	normal(0,38.91)
(7)	ψ_1	\overline{TA}_{my}	normal(0,4.9)
(7)	ψ_2	\overline{DL}_{my}	normal(0,0.62)
(7)	ψ_3	\overline{C}_{my}	normal(0,0.09)
(5), (6), (7)	Θ		exponential(1)

Table S3. Prior distributions of the regression coefficients and reciprocal dispersion parameter of the multilevel negative binomial models.