


ARTICLE

# Who Benefits? How Local Ethnic Demography Shapes Political Favoritism in Africa

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(Received 24 May 2019; revised 18 March 2020; accepted 16 April 2020; first published online 5 August 2020)

## Abstract

Empirical studies show that many governments gear the provision of goods and services towards their ethnic peers. This article investigates governments' strategies to provide ethnic favors in Africa. Recent studies of ethnic favoritism find that presidents' ethnic peers and home regions enjoy advantages, yet cannot disentangle whether goods are provided to entire regions or co-ethnic individuals. This article argues that local ethnic demography determines whether governments provide non-excludable public goods or more narrowly targeted handouts. Where government co-ethnics are in the majority, public goods benefit all locals regardless of their ethnic identity. Outside of these strongholds, incumbents pursue discriminatory strategies and only their co-ethnics gain from favoritism. Using fine-grained geographic data on ethnic demographics, the study finds support for the argument's implications in the local incidence of infant mortality. These findings have important implications for theories of distributive politics and conflict in multi-ethnic societies.

**Keywords:** ethnic politics; African politics; distributional conflict

A large body of research shows that where ethnicity is a salient political cleavage, governments often gear the provision of goods and services towards their ethnic peers (Burgess et al. 2015; Dreher et al. 2019; Franck and Rainer 2012; Hodler and Raschky 2014; Jablonski 2014; Kimenyi 2006; Kramon and Posner 2016). Ethnically biased goods provision slows economic development and can lead to distributional, sometimes violent ethnic conflict (Cederman, Weidmann and Gleditsch 2011; Easterly and Levine 1997). We build on studies demonstrating the existence of ethnic favoritism but move beyond them by investigating competing strategies of goods provision. We argue that governments can either provide regional public goods or target their ethnic peers with local club goods and private handouts. Understanding where governments provide which type of good yields important insights into the degree of ethnic discrimination that exists at various geographic levels.

Prior studies have generally operationalized government co-ethnicity either at the level of individuals or subnational geographic units. In doing so, they remain largely agnostic as to whether governments target co-ethnic individuals (for example, Franck and Rainer 2012) or co-ethnic regions (for example, De Luca et al. 2018; Hodler and Raschky 2014). This distinction is particularly relevant in Africa, where citizens' ethnic identity correlates with their place of residence, but not perfectly so. Thus individual and regional ethnic favoritism become observationally equivalent if tested using conventional research designs. Distinguishing the two strategies requires

information on government co-ethnics in largely non-co-ethnic areas and vice versa. If targeting is purely individual, citizens only benefit when their ethnic peers are in power, while if ethnic favoritism follows a regional logic, government co-ethnics and non-co-ethnics benefit in co-ethnic strongholds but not elsewhere.

In this article, we develop and test the more nuanced argument that local ethnic demography affects national governments' choices between broad regional and more precisely targeted provision strategies. In areas that are mainly populated by government non-co-ethnics, incumbent coalitions specifically target co-ethnic sublocalities and individuals. Since such highly localized targeting is comparatively expensive, governments provide locally non-excludable public goods in co-ethnic majority regions. These goods benefit all residents of a region, including non-co-ethnic minorities.

We use rich geographic data on local ethnic demographics and infant mortality to test the implications of our theory across twenty-two African countries. Infant mortality is a broad 'catch-all' proxy for the consequences of favoritism as it is affected by goods provided to individuals (for example, jobs and rents), villages (such as health clinics) or entire regions (for example, large-scale infrastructure). The Demographic and Health Surveys (DHS, USAID, 2012) contain information on the survival and death of more than 1.5 million African infants born between 1960 and 2013. Data on mothers' ethnic identities allow us to code individual co-ethnicity with the ruling coalition at the time of each infant's birth. The Spatially Interpolated Data on Ethnicity (SIDE, Müller-Crepon and Hunziker 2018) dataset provides data on local ethnic demographics that we use to construct the yearly, district-level population share that is co-ethnic with the national government. We use information on the ethnic composition of governments from the Ethnic Power Relations (EPR) data (Vogt et al. 2015).

With these data, we estimate the effects of changing co-ethnicity with the government at the individual and district levels. We identify these effects by only exploiting temporal variation in government co-ethnicity within ethnic groups and districts. Our findings support the implications of our argument on governments' differential goods provision strategies. We find that being born to a mother who is ethnically represented in the national government increases an infant's chance of survival by about 1.5 percentage points. However, this effect only holds in districts where the governing ethnic groups are in the minority at the time of birth, suggesting the provision of individually targeted benefits to government co-ethnics. Being born in a district with a high percentage of government co-ethnics similarly increases infant survival rates by 0.9–1.8 percentage points. In these districts, however, individual co-ethnicity with the government does not increase infant survival rates any further. This is consistent with the provision of locally non-excludable public goods.

Several robustness checks show that our results are not due to reverse causality, systematic migration patterns or the segregation of ethnic groups within districts. Measuring government co-ethnicity with data on individual cabinet members' ethnic identities from Francois, Rainer and Trebbi (2015) does not substantively alter our results. Additional analyses show that neither regime type nor electoral rules moderate our findings. Correlational evidence from Afrobarometer data indicates that governments indeed provide public goods to co-ethnic districts and more narrowly target their peers elsewhere.

## Literature

Ethnic favoritism constitutes the main material underpinning of political competition along ethnic cleavages. According to Bates (1974, 152), ethnic groups can be understood as coalitions that are formed to secure the scarce benefits of modernization and economic development. Ethnic markers such as individuals' language or phenotype enable an ethnically biased targeting of 'pork' (Fearon 1999) and thus foster the formation of ethno-political coalitions that benefit their members at the expense of other groups (Padró i Miquel 2007; Wantchekon 2003).

Empirical studies provide substantive evidence that governments favor their co-ethnic citizens and home regions (Burgess et al. 2015; Franck and Rainer 2012; Hodler and Raschky 2014). One strand of research investigates whether presidents provide benefits to individuals who share their ethnic identity. Franck and Rainer (2012) find that presidents' co-ethnics are better educated and enjoy lower infant mortality rates than other citizens. Kramon and Posner (2013), however, show that the effect of co-ethnicity with the president varies across countries and specific goods.

Another body of work operationalizes ethnic favoritism as an (ethno-)regional phenomenon. Burgess et al. (2015) report that Kenyan road investments during autocratic spells favor districts that share the president's ethnicity. Hodler and Raschky (2014) and De Luca et al. (2018) show that presidents' home regions experience faster growth measured by nightlights. Similarly, Dreher et al. (2019) find favoritism in Chinese-funded aid projects.<sup>1</sup> Kramon and Posner (2016) provide evidence that alternative measures of individual- and district-level co-ethnicity with the president both contribute to educational attainment in Kenya.

While sufficient to demonstrate the existence of favoritism per se, these two empirical approaches do not distinguish ethnic from geographic drivers of favoritism. Subnational regions in multi-ethnic countries are rarely perfectly homogeneous, and many citizens live outside their ethnic home region.<sup>2</sup> The empirical inability to differentiate the forms of ethnic favoritism comes with theoretical imprecision. Do governments target individuals or regions? Or do they follow a mixed strategy to minimize costs in the face of regionally varying ethnic demography?

Recent work on voting and distributive politics has started to disentangle the interplay between individuals' identities and local ethnic demography in Africa. In Ghana, Ichino and Nathan (2013) find that rural voters from local minority groups support non-co-ethnic majority candidates who promise locally non-excludable goods that benefit all residents. Similarly, Carlson (2015) and Nathan (2016) find that where one ethnic group dominates, residents expect public goods from politicians aligned with the local majority. However, local minority voters expect private goods from co-ethnic governments. Voters thus believe that politicians distribute goods in response to local demography. However, neither study tests whether actual goods provision reflects these patterns.

Two recent studies on local public goods provision provide evidence that local demography affects distributive politics. Ejdemyr, Kramon and Robinson (2018) argue that providing local club goods can only serve as a favoritism device if a politician's co-ethnics are geographically segregated. Data from Malawi supports their argument: ethnically segregated districts receive more boreholes, and within these districts, Members of Parliament (MPs) target mainly co-ethnic localities. Harris and Posner (2019) find similarly fine-tuned targeting of development projects in Kenya, where MPs tend to favor their co-ethnics and spatially segregated political supporters.

This article joins these studies in highlighting the importance of local ethnic demography but moves beyond them in four ways. First, our argument theorizes about the role of ethnic demography across a broader range of government-provided goods than very locally targeted club goods. Secondly, we show that not only geographically segregated co-ethnic strongholds but also co-ethnic minorities in 'opposition' districts profit from having their peers in power. Thirdly, we test whether non-co-ethnic individuals indeed benefit from living in co-ethnic strongholds, as argued by Ichino and Nathan (2013) and Ejdemyr, Kramon and Robinson (2018), and distinguish this effect from that of local ethnic segregation. Fourthly, covering twenty-two countries over 54 years, our analysis allows for more general conclusions than studies based on single countries, specific goods or cross-sectional variation.

<sup>1</sup>See also Jablonski (2014).

<sup>2</sup>The DHS data (USAID 2012) suggests that 84 per cent of the districts and 72 per cent of the enumeration areas in our sample are populated by more than one ethnic group.

## Theoretical Argument

We argue that areas where the majority population shares the national executive's ethnicity receive locally non-excludable club goods that benefit all residents. In areas with a small share of co-ethnics, governments provide more narrowly targeted goods to the few inhabitants who share their identity. As a result, government co-ethnics benefit from favoritism everywhere, whereas non-co-ethnics only benefit if they reside in governments' ethnic strongholds.

These predictions rest on three assumptions. First, the executive ruling coalition holds most of the power over distributive spending. Secondly, governments can more easily secure political support from co-ethnic citizens. Thirdly, the per-beneficiary provision cost of goods increases the more narrowly they are targeted to individuals. We start by highlighting the government branches that can best engage in ethnic favoritism, discuss why they would do so, and how they tailor goods provision strategies to local ethnic demography.

### Why do Co-Ethnics Benefit?

#### *Power and Incentives within Ruling Coalitions*

African state leaders want to stay in power but cannot rule alone. To survive politically, autocratic and democratic leaders need sufficient popular support to ward off revolutions, rebellions or electoral defeat (Bormann 2019; Francois, Rainer and Trebbi 2015; Svoblik 2012).<sup>3</sup> Leaders secure this support by targeting benefits to important constituencies.<sup>4</sup> Support buying is ethnically biased if in-group support is cheaper to buy than out-group backing. However, relying exclusively on co-ethnic support is often insufficient. As a remedy, most African leaders include elites from other ethnic groups in their cabinets (Bormann 2019; Francois, Rainer and Trebbi 2015).<sup>5</sup> To make such ethnic power sharing work, leaders need to give their coalition partners sufficient power and resources. Otherwise, their partners lose support from their constituencies (Beiser-McGrath and Metternich 2020), or worse, may threaten the regime from within (Roessler 2011). Potentially opposing their governments, African legislatures in electoral autocracies and (semi-)democratic regimes have long remained too weak to effectively constrain executive appointment and spending decisions (Fish and Kroenig 2009; Rainer and Trebbi 2014).<sup>6</sup> For these reasons and in line with past research, we view the executive branch of government, comprising the leader and her most important coalition partners, as the main locus of materially consequential power brokering (Arriola 2009; Rainer and Trebbi 2014).

#### *Buying Support from Co-ethnics*

Executive ruling coalitions prefer to buy support from their ethnic peers for several reasons. Citizens who derive expressive or 'psychic' utility from supporting co-ethnic candidates require fewer material incentives to be swayed (Chandra 2004). As such, they constitute a governing coalition's core support group and are the first to be targeted with goods and services (Cox and McCubbins 1986). Even where citizens are *a priori* indifferent between politicians from different groups and expected material benefits are all that matters, governments may serve their co-ethnics first. A common culture, language, dense social networks and higher trust in co-ethnics lower transaction costs, facilitate information flows and ease in-group policing (Fearon and Laitin 1996; Larson and Lewis 2017; Robinson 2020). Better information about co-ethnics' preferences and reliable local intermediaries enable governments to more efficiently

<sup>3</sup>The need for national-level support mainly holds in winner-takes-all executive elections rather than majoritarian legislative elections, where politicians face incentives to target contested swing districts (Casey 2015).

<sup>4</sup>See Golden and Min (2013) and Cox (2009) for useful reviews of the distributive politics literature.

<sup>5</sup>According to the EPR data set, 74 per cent of all country-years analyzed below were multi-ethnic. Francois, Rainer and Trebbi (2015) list not a single mono-ethnic cabinet.

<sup>6</sup>Note, however, that this changes as parties and legislatures become more important in consolidating democracies (Ariotti and Golder 2018; Barkan 2008; Opalo 2019).

allocate benefits (Baldwin 2016; Dixit and Londregan 1996). Similarly, promises of favoritism and political support are more credible coming from co-ethnics (Carlson 2015).

For governments, each unit spent on co-ethnics thus promises higher and more certain political support. While voters can of course also support non-co-ethnic candidates (Ichino and Nathan 2013; Nathan 2016), empirical evidence shows that good performance and clientelistic appeals by politicians generate more support among co-ethnic than non-co-ethnic voters (Adida et al. 2017; Carlson 2015; Kramon 2017; Wantchekon 2003). This evidence supports our assumption that buying political backing from ethnic peers is more cost effective for governments than buying it from other groups. Case studies suggest that this argument also holds in autocracies such as Nigeria under military rule. The ruling coalition of Hausa-Fulani and other Muslim elites from the North favored their ethnic and regional peers through redistricting reforms, federal revenue sharing formulas and rampant patronage (Abubakar 2001; Bah 2004).

### *Who Gets What?*

Governments have to decide where to provide what kinds of favors. We argue that differential provision costs for different types of goods as well as local ethnic demography inform this decision.

#### *Precise Ethnic Targeting is Costly*

Government-provided goods can be ordered along a continuum between local public goods and individually targeted private goods. At one end, there are non-excludable (local) public services such as schools and hospitals, or infrastructure such as highways that benefit all local residents regardless of their ethnic identity. At the other end of the spectrum, governments serve individually selected citizens with targeted handouts such as public employment, food and ultimately money (Albertus 2013; Besley et al. 2004). In between these extremes are local public goods that are targeted at ever smaller localities, for example, a small village road or a neighborhood water well. Goods thus vary in the degree to which their provision can exclude (groups of) citizens that are not intended to benefit.

We assume that the per-beneficiary cost of providing goods increases with their excludability. Locally non-excludable goods and infrastructures require high initial investments, but usage costs typically remain low. Once set up, a large number of local beneficiaries profits over extended periods. As a result, the provision cost per user is lower than for more narrowly targeted club goods and private handouts. The latter require more information on the intended recipients – locations, groups and individuals – and favors need to be customized to recipients' specific needs (Armesto 2010; Ichino and Nathan 2013).<sup>7</sup> While the non-discriminatory provision of public goods is cheaper on a per capita basis, governments have to pay the cost of local public services for every individual in the targeted locality, including citizens that the government does not intend to target (Nathan 2016). Thus, the cost efficiency of locally non-excludable goods increases with the number of recipients the government wants to reach (Ejdemyr, Kramon and Robinson 2018).

#### *Local Ethnic Demography and Goods Provision*

In short, governments face two, potentially conflicting, logics of goods provision. First, they prefer to target co-ethnic individuals. Secondly, governments prefer to deliver locally non-excludable goods that come at a lower per capita cost and benefit all citizens in their vicinity. Local ethnic population shares determine the optimal provision strategy that addresses this trade-off.

Homogenous co-ethnic areas constitute the easiest case. Here, governments want to reach almost all residents and therefore provide locally non-excludable goods. The dominant majority population in these government strongholds and the small local minority of non-co-ethnics benefit equally. Targeting co-ethnics in non-co-ethnic 'opposition' districts requires more expensive

<sup>7</sup>Governments may also transform local public goods into private ones by restricting access, which is costly.

club and private goods. Strategic incumbents only incur these costs if they rely on the support of co-ethnics in such non-co-ethnic areas. Where public goods provision to mainly co-ethnic regions ensures political survival, co-ethnics elsewhere may end up getting nothing and favoritism may turn into a purely regional phenomenon. However, if governments require at least some support from co-ethnics in districts where they are in the minority, they will serve them with targeted goods to save costs.

Ruling coalitions with a relatively small ethnic support base may additionally need the support of at least some non-co-ethnics. To minimize costs, they will provide public goods to areas with the highest proportion of co-ethnics first, before turning to non-co-ethnics in other areas. Similarly, the higher the costs of providing targeted benefits to co-ethnic individuals, the more governments will rely on public goods spending in increasingly mixed districts to sway non-co-ethnic minorities. As long as co-ethnics respond more readily to favoritism, however, targeted strategies remain rational in districts with low proportions of government co-ethnics.

Taken together, we argue that governments allocate non-excludable goods to districts with a high share of government co-ethnics. In areas dominated by ethnic groups not represented in the executive, governments target their co-ethnics with local club goods or private handouts. These two expectations lead to the following hypothesis:

**Hypothesis 1:** In districts with low shares of government co-ethnics, co-ethnic individuals are more likely to receive benefits from the government than non-co-ethnic individuals (*individual ethnic favoritism*). Districts with a higher share of co-ethnics receive greater benefits, but an inhabitant's ethnic identity has no effect on her likelihood of receiving goods from the government (*regional ethnic favoritism*).

## Data

We test our theory with a large, individual-level dataset on infant mortality covering twenty-two African countries and 54 years (1960–2013). We expect our argument to be relevant in other multi-ethnic sub-Saharan African cases as well, but the availability of geocoded data on our main outcome measure, infant mortality, limits our sample.<sup>8</sup>

## Outcome Data

Our infant mortality data come from the DHS (USAID, 2012). Female DHS respondents' birth histories record the survival and death of 1.5 million infants born between 1960 and 2013.<sup>9</sup> The surveys include mothers' ethnic identity, current location, and their children's birth years. This information and the large sample size allow us to exploit cross-sectional variation in mortality rates within countries and temporal variation within districts and ethnic groups.

The usefulness of infant mortality as a proxy for both individual and regional favoritism depends on the extent to which mortality rates respond to government-provided private and public goods. Both relationships are well documented in the literature. Private goods such as public sector jobs (Posner 2005), access to local markets (Bates 1974), or handouts such as medication or other supplies<sup>10</sup> can affect infant mortality either directly (as in the case of medical supplies) or by increasing household income. Increases in household income have been shown to mitigate various risk factors such as malnutrition, respiratory infections and malaria among vulnerable children in the Democratic Republic of the Congo (Grellety et al. 2017), Burkina Faso

<sup>8</sup>Although ethno-regional favoritism is not unique to Africa (De Luca et al. 2018), the validity of our argument beyond our sample depends on whether our main assumptions hold in the respective cases.

<sup>9</sup>We drop infants born less than 12 months before an interview.

<sup>10</sup>E.g., fertilizer as studied by Abman and Carney (2019).



(Houngbe et al. 2017) and Kenya (Huang et al. 2017). Similarly, conditional cash transfers combined with health or nutritional programs decrease child mortality (Barham 2011; Rasella et al. 2013).

With respect to public goods, previous research highlights the importance of infrastructure (Fay et al. 2005), especially public health care facilities (Gruber, Hendren and Townsend 2014) and sanitation systems (Galiani, Gertler and Scharrodsky 2005). Of course, other types of public goods such as roads may also affect infant mortality via their effects on household income and better access to health facilities.

While infant mortality rates plausibly capture various types of government-provided goods, they do not allow us to distinguish excludable private from non-excludable public goods directly. We address this shortcoming by (1) testing whether individual-level ethnic favoritism is apparent within DHS survey clusters (typically villages or urban neighborhoods) and (2) analyzing cross-sectional data on local public services provision from the Afrobarometer surveys.

### Co-ethnicity with the Government

To test for individual-level favoritism, we measure infants' co-ethnicity with the government by linking their mothers' ethnic identities to the ruling ethnic groups in the respective country and birth year. Information on ethnic representation in executive governments comes from the EPR dataset (Vogt et al. 2015).<sup>11</sup> For each country-year, the EPR lists all politically relevant ethnic groups<sup>12</sup> and codes whether they have meaningful representation in the national executive as opposed to mere window dressing. Aiming to identify ethnic groups whose elites hold power and control distributive resources in the executive, the data thus code multi-ethnic coalitions, which are frequent in Africa (Francois, Rainer and Trebbi 2015).<sup>13</sup> If government elites from other than the president's group indeed engage in favoritism, this coalition-based approach is more precise than a mere focus on heads of state.<sup>14</sup>

To operationalize the regional aspect of favoritism, we measure the proportion of each district's population that is co-ethnic with the government in a given year. The SIDE dataset on local ethnic demographics in Africa provides non-parametric spatial interpolations of DHS survey locations' ethnic population shares (Müller-Crepon and Hunziker 2018). For each DHS survey, SIDE predicts cell-level ethnic compositions on a raster with about  $1 \times 1$  km resolution. Although the interpolation increases precision over the raw DHS data or more commonly used polygon data, the estimates are not precise enough to capture local segregation in ethnically mixed areas.<sup>15</sup>

Linking the SIDE data to EPR and aggregating it to the district level, we calculate the percentage of government co-ethnics for each district, year and DHS survey round as:

$$\text{District Share Co-Ethnic}_{dts} = \frac{\sum_{c=0}^{C_d} \sum_{e=0}^E \text{pop}_{ct}^* \text{share}_{ecs}^* \text{incl}_{et}}{\sum_{c=0}^{C_d} \text{pop}_{ct}},$$

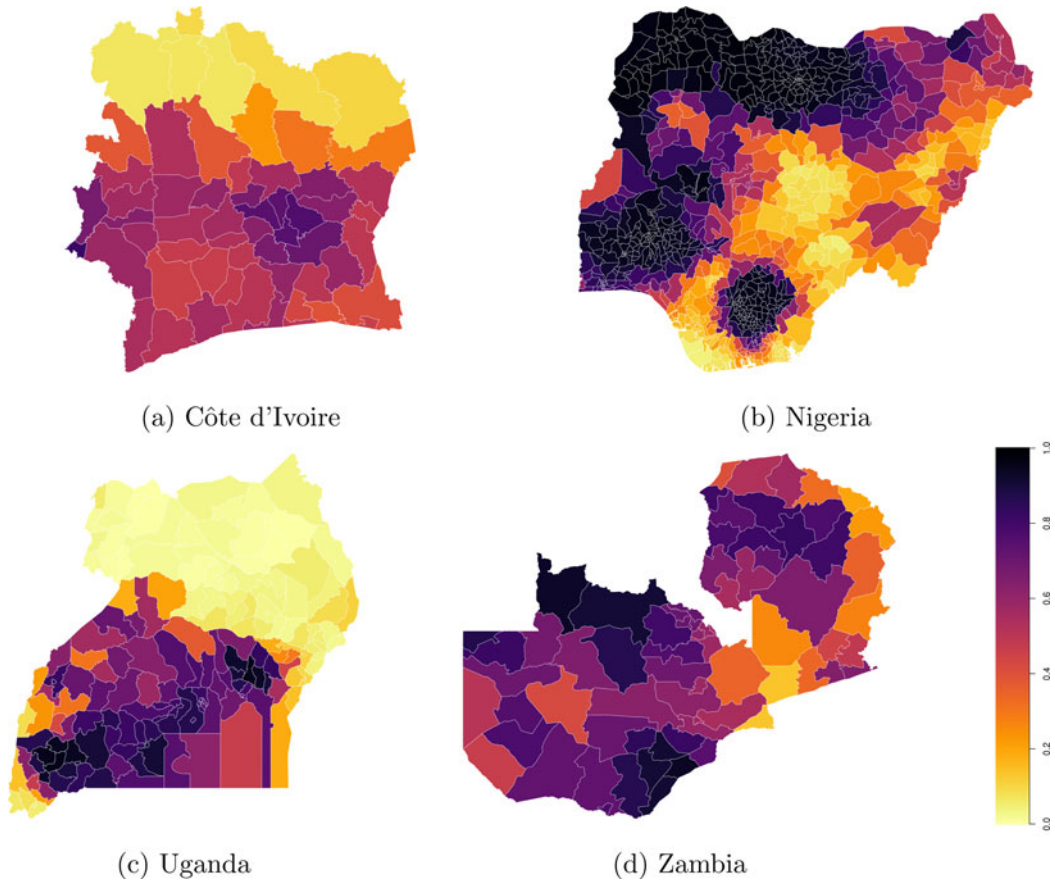
<sup>11</sup>See Figure A5 for the ethnic power constellations in each country-year. Groups are matched based on group names and information from encyclopedias such as ethnologue.com, wikipedia.com, and joshuaproject.org.

<sup>12</sup>The EPR codes ethnic groups as politically relevant if they are politically mobilized at the national level or discriminated against by the government (Vogt et al. 2015).

<sup>13</sup>The EPR provides an ordinal measure of ethnic representation and does not list the positions and elites informing this proxy. Francois, Rainer and Trebbi (2015) provide alternative data on the ethnic composition of cabinets in fifteen countries, used in a robustness check in Appendix Table A11.

<sup>14</sup>But see the robustness check in Appendix Table A7, in which we analyze EPR's 'senior partners', in most cases the presidents' ethnic groups.

<sup>15</sup>Since no non-ethnic covariates affect the predictions, interpolation errors constitute random noise.



**Figure 1.** District-level co-ethnicity with the government in 2000

Note: figure uses the most recent available SIDE data. See Appendix Figure A3 for maps of all other countries in our sample.

where  $d$  denotes a district in year  $t$  with SIDE information on ethnic groups  $E$  estimated on the basis of DHS survey  $s$ .  $C_d$  are all raster cells  $c$  in district  $d$ ,<sup>16</sup> and each cell has a population of  $\text{pop}_{ct}$ .<sup>17</sup> In each raster cell, each ethnic group  $e$  has an estimated population share  $\text{share}_{ecs}$  derived from the SIDE data for DHS surveys. Linking ethnic group  $e$  to EPR provides information on its representation in government:  $\text{incl}_{et} \in [0;1]$ . In sum, this spatial computation yields, for each district-year per survey, the population share of government co-ethnics. Our measure exhibits substantial spatial variation within countries (Figure 1).<sup>18</sup> Similarly, variation in the ethnic composition of governments produces large changes within districts over time (Figure 2).

Finally, we assign District Share Co – Ethnic<sub>dt</sub>s to infants via mothers' geocoded enumeration area, survey round  $s$ , and each infant's year of birth  $t$ . Calculating district-level co-ethnicity on a per-survey basis ensures that the coding of ethnic groups, which changes between DHS rounds, remains consistent when we match an infant to the district-level data.<sup>19</sup>

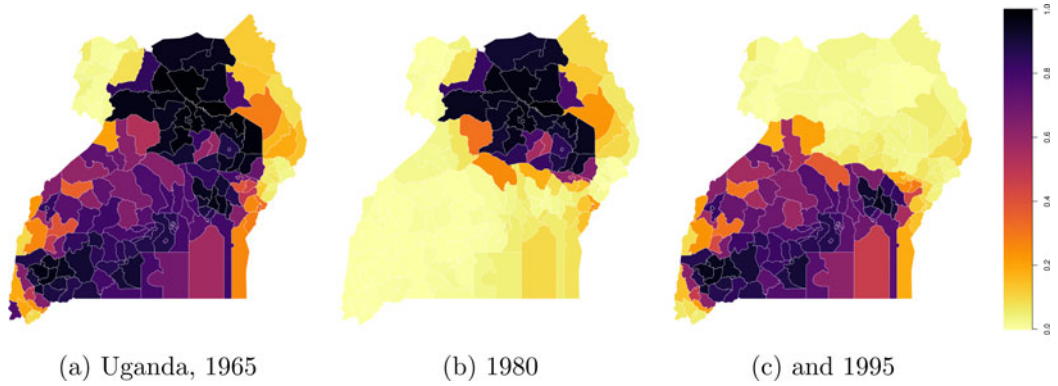
<sup>16</sup>District borders are from 2000 and come from FAO (2014).

<sup>17</sup>Population rasters with the same resolution as SIDE come from CIESIN et al. (2011) and cover the years 1990, 1995 and 2000. We take the latest year available for each  $t$  and assume that changes in districts' population distribution only marginally affect the ethnic compositions.

<sup>18</sup>Co-ethnicity shares are bimodally distributed, with 60 per cent of districts with shares below 10 per cent or above 90 per cent, and 40 per cent in between; 33 per cent of DHS enumeration area-years have 10–90 per cent government co-ethnics.

<sup>19</sup>Appendix Table A2 shows that both measures coincide.





**Figure 2.** District-level co-ethnicity with the government over time  
Source: SIDE data from 2011.

Since we lack data on migration patterns, we take each mother's place of residence at the time of the survey as all her children's birthplace. In parallel, we assign each district a constant ethnic demography. To identify valid effects, we therefore assume that, conditional on several fixed effects, unobserved migration is orthogonal to spatio-temporal changes in ethnic inclusion and infant mortality. We address this assumption below.

### Empirical Strategy

Our models estimate the effect of individual- and district-level co-ethnicity with the government on infant mortality. To test our intuition that governments only discriminate between their co-ethnics and other citizens in districts where the co-ethnic share is low, we include an interaction term that assesses whether district-level co-ethnicity with the government moderates any individual-level co-ethnicity advantage. We use linear probability models to enable a fixed-effects strategy that controls for confounding characteristics of ethnic groups, districts and temporal shocks.<sup>20</sup>

Our baseline specification takes the following form:

$$\begin{aligned}
 Y_{ierst} = & \alpha_{es} + \lambda_{ds} + \gamma_{rst} + \beta_1 \text{Co-Ethnic Government}_{est-1} \\
 & + \beta_2 \text{District Share Co-Ethnic}_{dst-1} + \beta_3 \text{Co-Ethnic Government}_{est-1} \\
 & \times \text{District Share Co-Ethnic}_{dst-1} + \delta X_{ierst} + \epsilon_{ierst}.
 \end{aligned}$$

The unit of analysis is infant  $i$  born in year  $t$  to a mother who hails from ethnic group  $e$ , resides in district  $d$  in subnational region  $r$ , and was interviewed in DHS country-survey round  $s$ .  $Y_{ierst}$  is a dummy variable coding the death ( $Y=100$ ) or survival ( $Y=0$ ) of infant  $i$  in the first twelve months after its birth. We code the dummy as 100 to straightforwardly interpret the coefficients in terms of percentage points.

The main variables of interest are a mother's co-ethnicity with the government, the district-level population share of government co-ethnics, and the interaction term of these two variables.<sup>21</sup> We temporally lag these predictors by one year to allow the hypothesized effects to unfold.<sup>22</sup> Our

<sup>20</sup>Absent well-founded assumptions about the 'true' functional form, the linear probability model does not perform worse than logit or probit models (Angrist and Pischke 2008).

<sup>21</sup>See, e.g., Sands and de Kadt (2019) for a similar but not equivalent modelling strategy.

<sup>22</sup>Our results are robust to using alternative time lags (Appendix Table A3).

hypothesis stipulates that  $\beta_1$ , the coefficient of individual-level co-ethnicity, is negative because governments favor their co-ethnics everywhere. Co-ethnics in government strongholds receive local public goods, whereas co-ethnic minorities in opposition districts are targeted individually. As public goods also benefit non-co-ethnic residents, we expect  $\beta_2$  to be negative. Finally, because co-ethnics in mainly co-ethnic districts receive the same public goods as non-co-ethnics and no additional private handouts, we expect the interaction effect  $\beta_3$  to be positive.

In order to credibly identify the effects of our three variables of interest, the model only exploits yearly variation in ethnic representation as coded in the EPR data. We treat these changes as shocks that differentially affect infants born in the same year and to mothers from the same ethnic group and district. We thus add three sets of fixed effects.

The first set,  $\alpha_{es}$ , is a vector of ethnic-group-survey-round fixed effects. They ensure that we only compare infants within ethnic groups whose mothers were interviewed in the same survey. This controls for time-invariant characteristics of ethnic groups such as their demographic size or geographic settlement area that may simultaneously affect infant mortality and their political representation. We use separate fixed effects for each survey round to ensure that mothers' ethnic identities line up with the SIDE district shares derived from the same DHS surveys.

Secondly,  $\lambda_{ds}$  is a vector of district-survey-round fixed effects that controls for unobserved, time-invariant differences between districts such as resource endowments. These ethnic group and district fixed effects imply that we only identify effects from temporal variation within ethnic groups and districts.

Thirdly, region-survey-birthyear fixed effects  $\gamma_{rst}$  flexibly control for temporal shocks to infant mortality that differentially affect first-level administrative units. Subnational shocks such as differential economic growth may correlate with government composition and bias results from models with looser temporal controls. In fact, infant mortality rates in sub-Saharan Africa exhibit far greater within-country than between-country variation (Burke, Heft-Neal and Bendavid 2016). These region-year fixed effects also absorb all common shocks that operate at the continent or country level, for example institutional reforms or changes in government budgets.

$X_{ierst}$  is a vector of individual control variables including female and twin dummies, an infant's birth rank and its square, as well as mothers' age and age square at the time of birth. We report two-way clustered standard errors at the level of both ethnic and district-survey rounds to account for cross-sectional and serial correlation across infants and birth cohorts within the units affected by our treatment.<sup>23</sup> Regression weights ensure that the twenty-two countries in our sample contribute equally to the estimated effects.<sup>24</sup>

## Results

Our baseline specification as well as increasingly conservative models yield support for our hypothesis. In districts with high shares of government co-ethnics, infants from *all* ethnic groups exhibit lower mortality rates. In districts with few government co-ethnics, however, this is only the case for co-ethnic children.

Model 1 in Table 1 presents the results of our main specifications. The coefficient of the individual-level government co-ethnicity variable is negative and statistically significant. It suggests that co-ethnic infants born in districts with very low shares of government co-ethnics have a 1.49-percentage-point lower mortality risk than non-co-ethnic infants in these districts. The coefficient on district-level government co-ethnicity is also significantly negative. Increasing the share of co-ethnics in an infant's district by 50 percentage points is associated with a decrease in non-co-ethnic mortality by 0.91 percentage points.

<sup>23</sup>Our results are robust to clustering at different resolutions (see Appendix Table A4).

<sup>24</sup>Weights are computed as the inverse of the number of observations from each country. They avoid undue influence of countries that are (a) more frequently surveyed (i.e., more developed) by the DHS or (b) have high fertility rates (i.e., less developed). Unweighted regressions lead to similar results (Appendix Table A5).

Table 1. Main specifications

	Infant mortality			
	1	2	3	4
Government co-ethnic ( $t-1$ )	-1.487*** (0.391)		-1.651*** (0.462)	
Dist. share gov. co-ethnics ( $t-1$ )	-1.816*** (0.637)	-1.687* (0.878)		
Co-ethnic $\times$ dist. share co-ethnics ( $t-1$ )	1.824*** (0.503)	1.880*** (0.554)	1.835*** (0.537)	1.714*** (0.582)
Survey-ethnic FE	Yes	-	Yes	-
Survey-district FE	Yes	Yes	-	-
Survey-region-birthyear FE	Yes	Yes	-	-
Survey-ethnic-birthyear FE	No	Yes	No	Yes
Survey-district-birthyear FE	No	No	Yes	Yes
Controls	Yes	Yes	Yes	Yes
Observations	1,503,930	1,503,930	1,503,930	1,503,930
Adjusted $R^2$	0.060	0.061	0.068	0.069

Notes: OLS linear probability models. The sample mean of the dependent variable is 10.78 infant deaths per 100 live births. Observations are weighted to ensure equal weights for each country. Control variables include mothers' age and age squared as well as infants' sex, a twin dummy, birth rank and birth rank squared. Two-way clustered standard errors in parentheses (survey-ethnic group and survey-district clusters). \* $p < 0.1$ ; \*\* $p < 0.05$ ; \*\*\* $p < 0.01$

The coefficient of the interaction term of individual-level co-ethnicity  $\times$  district-level share of co-ethnics is significantly positive and about the same size as the constitutive terms. Figure 3 visualizes how the interplay of individually and regionally targeted ethnic favoritism affects infants' mortality. A non-co-ethnic infant in a district with no government co-ethnics serves as the baseline category to which we compare predicted outcomes for infants in all other categories.<sup>25</sup> The x-axis displays the proportion of non-co-ethnics residing in a district. At low district shares of government co-ethnics, co-ethnic infants (dashed line) have a 1.49-percentage-point lower mortality than their non-co-ethnic counterparts (solid line). As the share of government co-ethnics rises, co-ethnic infants' estimated survival advantage over non-co-ethnic infants in non-co-ethnic districts remains essentially constant. Mortality estimates for non-co-ethnic infants show a markedly different pattern. With increasing shares of government co-ethnics in a district, non-co-ethnic infants' mortality rates decrease. At very high district shares of government co-ethnics, non-co-ethnic infants' predicted mortality rate is about 1.82 percentage points below that of non-co-ethnic infants in entirely non-co-ethnic districts. In districts with high shares of co-ethnics, co-ethnic infants have no advantage over non-co-ethnic children.<sup>26</sup>

The size of these effects is substantial. The average infant mortality rate in our sample drops by about 0.2 percentage points each year.<sup>27</sup> Thus the disadvantage of non-co-ethnic children in non-co-ethnic regions compared to co-ethnic children or those born in co-ethnic districts amounts to between 7 and 9 years of that trend.

As our hypothesis suggests, co-ethnic and non-co-ethnic mortality rates only differ significantly in districts where government co-ethnics are the minority. Where government co-ethnics constitute more than about half of a district's population, co-ethnic and non-co-ethnic mortality rates are statistically indistinguishable. These findings are consistent with the provision of locally

<sup>25</sup>These predictions visualize the substitution effect between individual- and district-level co-ethnicity. Predictions for co-ethnics are calculated as  $\beta_1 + a_d \beta_2 + a_d \beta_3$ ; those for non-co-ethnics as  $a_d \beta_2$  where  $a_d$  is the district share of government co-ethnics in hypothetical district  $d$ . Appendix Figure A6 shows conventional marginal effect plots of individual co-ethnicity across the range of district-level government co-ethnicity.

<sup>26</sup>The binning estimator proposed by Hainmueller, Mummolo and Xu (2019) demonstrates that our findings are not an artifact of linearity assumptions or a lack of common support. See Appendix Figures A7 and A8.

<sup>27</sup>See Appendix Figure A2.

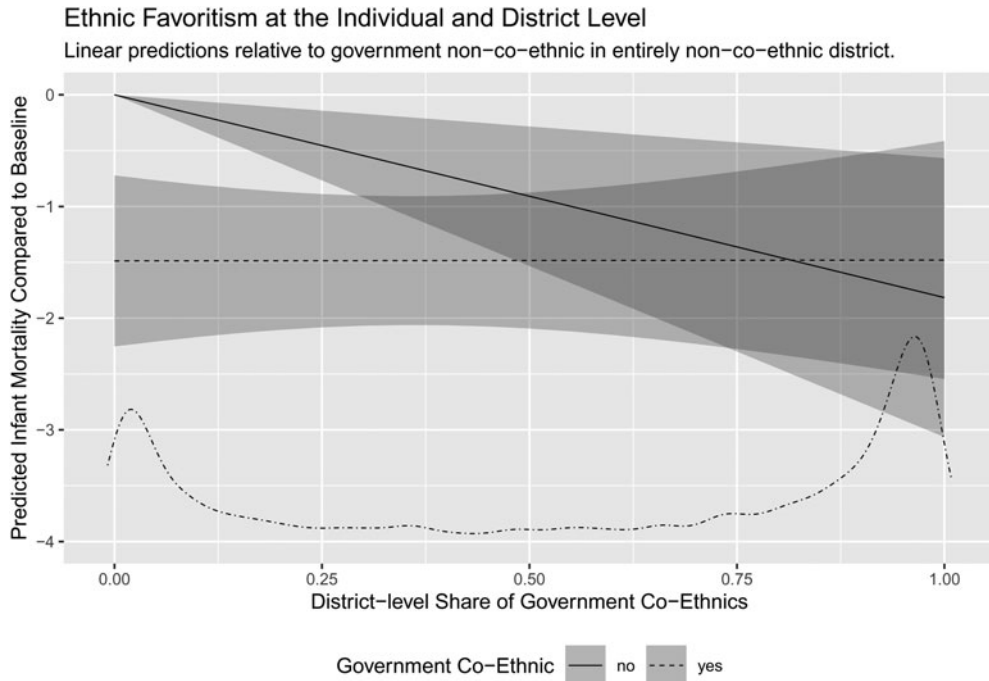


Figure 3. Predictions for government co-ethnics and non-co-ethnics conditional on local ethnic demography

non-excludable goods in ethnic government strongholds and the provision of excludable goods in government minority districts.

Models 2–4 in Table 1 add fixed effects that account for unobserved temporal shocks at the level of districts and ethnic groups. Both may affect the ethnic composition of governments and bias our estimates. Model 2 includes survey-ethnic-group-year-of-birth fixed effects. They control for all yearly events at the level of groups, such as changes in economic development or the outbreak of violence. These fixed effects fully absorb variation in individual-level government co-ethnicity; the respective constitutive term remains unidentified. Model 3 follows a similar logic and adds survey-district-year-of-birth fixed effects that control for shocks to a given district. Comparing only infants born in the same year and district, the model no longer identifies the district-level constitutive term. Finally, Model 4 combines both of these spatio-temporal fixed effects and thus only identifies the interaction term. Throughout these specifications, the identified coefficients of interest remain stable and statistically significant. Our results are thus not caused by time-variant omitted variables at the level of ethnic groups or districts.

### Alternative Explanations and Sensitivity Analyses

In this section, we provide evidence that our results are not driven by within-district ethnic segregation, non-parallel trends between ethnic groups or unobserved migration patterns. We then show that the results are robust to an alternative operationalization of co-ethnicity with the government. Lastly, our heterogeneity analysis fails to detect any significant moderating effects of regime type or electoral systems.

#### *Ethnic Segregation within Districts*

The effects of both individual- and district-level government co-ethnicity in the main analyses may not result from the provision of different types of goods but from uniform club goods

**Table 2.** Robustness: cluster fixed effects, trends, and subsample analysis

	Infant mortality			
	Cluster-YoB FE 1	Time trend 2	Pre-trends 3	Less educated 4
Government co-ethnic ( $t-1$ )	-1.506** (0.698)	-1.318*** (0.486)	-1.662*** (0.420)	-1.027** (0.436)
Dist. share government co-ethnics ( $t-1$ )		-1.835** (0.771)	-2.094*** (0.703)	-1.654** (0.699)
Co-ethnic $\times$ dist. share co-ethnics ( $t-1$ )	0.928 (0.776)	1.891*** (0.540)	1.941*** (0.517)	1.228** (0.594)
Upgrade $_t$ to $t+2$			-0.619 (0.503)	
Downgrade $_t$ to $t+2$			-0.248 (0.495)	
Survey-ethnic FE	Yes	Yes	Yes	Yes
Survey-district FE	-	Yes	Yes	Yes
Survey-region-birthyear FE	-	Yes	Yes	Yes
Survey-cluster-birthyear FE	Yes	No	No	No
Survey-ethnic time trend	No	Yes	No	No
Controls	Yes	Yes	Yes	Yes
Observations	1,503,930	1,503,930	1,483,404	1,315,931
Adjusted $R^2$	0.099	0.060	0.060	0.062

Notes: OLS linear probability models. Column 4 restricts the sample to children born to mothers with less than secondary education. The sample mean of the dependent variable is 10.77 in Columns 1–3 and 11.37 in Column 4. Observations are weighted to ensure equal weights for each country. Control variables include mothers' age and age squared, as well as children's sex, a twin dummy, birth rank, and birth rank squared. Two-way clustered standard errors in parentheses (survey-ethnic group and survey-district clusters). \* $p < 0.1$ ; \*\* $p < 0.05$ ; \*\*\* $p < 0.01$

provision to mainly co-ethnic localities in ethnically segregated districts (as in Ejdemyr, Kramon and Robinson 2018). Homogeneous co-ethnic villages rather than co-ethnic individuals might be the primary beneficiaries of ethnic favoritism in mainly non-co-ethnic districts. Similarly, the district-level effects might be driven by government non-co-ethnics who happen to reside in relatively homogeneous co-ethnic villages rather than by entirely non-co-ethnic villages benefiting from broader district-level public goods.

We estimate three additional models to address this inferential threat. First, we limit comparisons to government co-ethnics and non-co-ethnics born in the same year and survey cluster, typically a village or urban neighborhood. The results in Model 1 of Table 2 suggest that government co-ethnics born in survey clusters in non-co-ethnic districts have, on average, better survival chances than non-co-ethnic infants born in the same year and location.<sup>28</sup> Secondly, we define regional favoritism at the survey cluster rather than the district level. Even in mainly non-co-ethnic enumeration areas, co-ethnic infants survive longer (Model 2 in Appendix Table A10). Thirdly, we aggregate our main variables to the level of DHS enumeration-area-years. In districts with high shares of government co-ethnics, even entirely non-co-ethnic enumeration areas show significantly lower infant mortality (Model 3 in Table A10). The very local individual and broader district-level effects suggest that our baseline results are not entirely due to local club goods provided to co-ethnic villages in segregated districts.

**Reverse Causation and Non-parallel Trends**

To interpret our results as causal, we must assume that changes in the ethnic composition of ruling coalitions are exogenous to observed or correctly anticipated trends of ethnic groups' infant mortality rates and their causes. If, for example, ethnic groups' economic development affects

<sup>28</sup>Additional models in Appendix Table A14 limit comparisons to children born (1) within the same household and (2) to the same mother. While standard errors increase due to a drastic reduction in the effective sample size and many coefficients miss significance, the substantive pattern in our point estimates remains consistent with the baseline results.

their political inclusion, reverse causation can bias our findings. To address this potential caveat, we include dummies for the three years before an ethnic group gets upgraded to or excluded from the national executive.<sup>29</sup> The results from Model 3 in Table 2 do not show any evidence of differential infant mortality in these periods. Additional results from regressions including individual dummies for the three (one) year(s) prior to government changes or linear time trends for three or five years before such changes are reported in Appendix Table A9. All of these terms remain small and insignificant or point in the direction that supports our argument. This reduces our concerns about reverse causation, anticipation effects and non-parallel trends.

### Migration

Unobserved migration patterns and other demographic shifts present a third challenge to our findings. Recall that we use a mother's place of residence at the time of the DHS interview as the birthplace of her children, thereby introducing measurement error. Similarly, the ethnic group shares from SIDE are assumed to be constant. All temporal variation in the district share of government co-ethnics thus comes from changes in governments' ethnic composition. If political inclusion correlates with fertility rates or migration patterns, our measure of district-level co-ethnicity will be slightly off. These patterns bias our results if they co-vary with government changes in a way that is not captured by ethnic group- and district-year fixed effects and the pre-trends analyzed above.

However, well-off non-co-ethnics from poor districts may systematically migrate to areas that are strongholds of a *new* government in order to access emerging opportunities. Similarly, if the incoming government coalition hands out jobs to co-ethnics, newly appointed and comparatively wealthy state bureaucrats may be posted to non-co-ethnic districts. To bias our results, such migration patterns have to temporally co-vary with government changes and be distinct from the overall temporal trend at the level of districts and ethnic groups. We regard such scenarios as rather unlikely.

If relevant, such migration patterns most likely apply to relatively educated individuals. Model 4 in Table 2 therefore drops mothers with more than primary education. The relevant coefficient sizes are somewhat smaller than before but remain substantively and statistically significant.<sup>30</sup>

### Co-ethnicity with Whom?

As an alternative to the EPR data on ethnic power constellations, we use Francois, Rainer and Trebbi's (2015) data on cabinet members' ethnic identities. We replicate our baseline model for the subset of thirteen countries and 44 years for which both minister data and DHS birth records are available. The results remain similar to the EPR analysis, especially if we define our treatment as co-ethnicity with Francois, Rainer and Trebbi's (2015) 'top government' ministers (see Table A11). These results are consistent with our argument that not only state leaders but also powerful ethnic coalition partners have control over distributive spending.

### Alternative DHS Outcomes

We capture the provision and use of health services more directly through DHS data on prenatal and birth assistance, the use of birth clinics and newborns' weight. Unfortunately, while mothers report the survival and death of all their children, these items are only enumerated for children born within five years before each survey. The respective sample is therefore smaller and covers a much shorter time span than our baseline data. Nonetheless, both cross-sectional (Table A12)

<sup>29</sup>See Hodler and Raschky (2014) for a similar approach.

<sup>30</sup>A lower likelihood of less educated respondents to report infant deaths may also cause this divergence.



and within-district and ethnic group estimates (Table A13) reveal similar main and interactive effects of individual- and district-level co-ethnicity with the government as observed in the infant mortality specifications.

### **Heterogeneous Effects**

Finally, we investigate whether democratic institutions moderate our findings as they constrain executive governments.<sup>31</sup> The mixed results presented in Appendix A3 suggest that democratic institutions may reduce but not eliminate favoritism. An exploration of the impact of varying electoral systems on ethnic favoritism yields no statistical difference between proportional representation and first-past-the-post systems. These findings suggest that executive governing coalitions have incentives to favor their ethnic constituents even in the absence of democratic elections and regardless of specific electoral rules.

### **Evidence on Ethnic Favoritism from the Afrobarometer**

We argue that governments provide public goods in co-ethnic districts and more narrowly target co-ethnics in non-co-ethnic districts. While the main analysis tests the distributional consequences of the argument, our data on infant mortality are unable to distinguish between public and private goods provision. To explore this question in more detail, we turn to Afrobarometer (2015) survey data. These data contain information on governments' provision of public services, which allows us to test whether they provide non-excludable goods to co-ethnic districts. In addition, we analyze whether patterns in households' economic well-being coincide with our findings on infant mortality.

We link the geocoded Afrobarometer rounds 1 to 5 (Afrobarometer, 2015; Ben Yishay et al. 2017) to the EPR and SIDE data using the same procedures as described above (see also Appendix A4). We then estimate the effect of individual- and district-level government co-ethnicity on the reported ease of (1) accessing public services and (2) households' economic well-being based on questions about how often respondents have 'gone without' food, water, health care and income. Below, we report effects on principal components that summarize items in both categories, and those items that directly relate to health care and thus infant mortality.<sup>32</sup> Because the Afrobarometer data lacks the temporal depth necessary for difference-in-differences models, we only exploit cross-sectional information within survey rounds in the same country. Omitted variables at the ethnic group and/or district level might therefore bias the results below.

Model 1 in Table 3 shows that public services such as health care and primary schooling are more accessible to respondents in co-ethnic districts. However, there is no association between individual-level government co-ethnicity and public service access. This pattern also holds with regard to accessing medical services (Model 2) and supports the argument that co-ethnic districts receive more public goods.

Turning to households' well-being, Model 3 shows a negative association between government co-ethnicity at the individual and district levels and the economic hardship factor. In line with our main analysis, there is no distinguishable difference between co-ethnics and non-co-ethnics in districts with a majority of government co-ethnics. The same pattern affects the reported frequency with which respondents or their family members lack access to medicines or medical treatment (Model 4). These findings closely align with our main results.

In combination, the results suggest that while co-ethnics in non-co-ethnic areas may not have better access to *public services* than their neighbors, they are substantively better off than non-co-ethnics. This divergence is consistent with our argument since *private goods* such as subsidies and jobs provided by their government may explain the divergence between access to public services and households' access to food, water and health care.

<sup>31</sup>Previous studies provide mixed findings on the effect of democracy on favoritism (Burgess et al. 2015; Franck and Rainer 2012; Hodler and Raschky 2014; Kramon and Posner 2016).

<sup>32</sup>Appendix A4 reports disaggregated analyses.

**Table 3.** Economic hardship and public services: cross-sectional OLS

	Ease of public service access		Economic hardship	
	PC 1	Medical services (1–4) 2	PC 3	Medical treatment (0–4) 4
Government co-ethnic ( $t-1$ )	–0.100 (0.162)	0.056 (0.088)	–0.382*** (0.136)	–0.239*** (0.088)
Dist. share gov. co-ethnics ( $t-1$ )	0.284** (0.125)	0.165*** (0.063)	–0.745*** (0.117)	–0.445*** (0.072)
Co-ethnic $\times$ dist. share co-ethnics ( $t-1$ )	0.117 (0.211)	–0.076 (0.107)	0.504*** (0.168)	0.307*** (0.106)
Individual-level covariates:	Yes	Yes	Yes	Yes
Country-survey fixed effects:	Yes	Yes	Yes	Yes
Observations	26,418	38,950	64,026	70,432
Adjusted $R^2$	0.076	0.055	0.150	0.123

Notes: OLS linear models. Control variables include four levels of education, age and age squared, as well as a female dummy. Two-way clustered standard errors in parentheses (language group and district clusters). \* $p < 0.1$ ; \*\* $p < 0.05$ ; \*\*\* $p < 0.01$

## Conclusion

This article studies how local ethnic demography moderates patterns of ethnic favoritism in twenty-two African countries. We argue that governments try to favor their co-ethnics at minimal cost, tailoring the types of goods they provide to local ethnic population shares. Co-ethnic government strongholds receive locally non-excludable public goods that also benefit non-co-ethnic minorities. Where governments' co-ethnics are in the minority, by contrast, they discriminate and reach their peers with more narrowly targeted goods and services.

We test these predictions using data on infant mortality and district-level ethnic demographics in sub-Saharan Africa that allow us to disentangle the effects of individual- and district-level government co-ethnicity. Our results are consistent with the provision of (1) non-excludable public goods to ethnic government strongholds and (2) ethnically targeted local or individual-level handouts to co-ethnics in opposition districts. Infants born into an ethnic group in power have substantially higher chances of survival, regardless of where they live. Conversely, government non-co-ethnics only have comparably low mortality rates if they are born in districts in which governing groups constitute the local majority. These effects are identified based only on temporal variation within ethnic groups and districts, which makes it unlikely that they are caused by omitted variable bias or reverse causality.

The insight that ethnic geography affects strategies of favoritism has three broader implications for the study of distributive politics in multi-ethnic societies. First, we highlight the importance of local ethnic demography for understanding the interplay between the ethno-political macro and micro levels. Secondly, we build on previous literature and stress that governments choose between locally non-excludable public goods and individually targeted handouts when designing distributive policies. Unfortunately, and absent better data on public and private goods provision, we are not yet able to optimally operationalize these different strategies with cross-national data. Thirdly, and in line with Ichino and Nathan (2013), voters may adjust their voting behavior to governments' goods provision strategies. Non-segregated ethnic geographies and other incentives for the provision of public goods may thus reduce the political salience of ethnic identities and identity-based distributive conflicts.

The dire effects of ethnic favoritism on children's lives suggest that equitable development policy should focus on those who are disenfranchised from government services. Our findings can be used to identify which citizens are least likely to be served by the government based on the interplay between their geographic location and ethnic identity. Citizens who are not ethnically represented in government *and* reside in areas where the majority shares that fate are the least likely to benefit from government services. This information can help to target aid at the most vulnerable and alleviate ethnic inequality.

**Supplementary material.** Data replication sets are available in Harvard Dataverse at: <https://doi.org/10.7910/DVN/DJR0YC> and online appendices at: <https://doi.org/10.1017/S0007123420000241>.

**Acknowledgements.** We thank Sona Golder and three anonymous referees for excellent comments and suggestions. Previous versions of this article were presented at the European Political Science Association Annual Conference 2018, at the Annual Meeting of the section ‘Political economy’ of the German Political Science Association 2019 and at two workshops at the University of Konstanz. Janina Beiser-McGrath acknowledges support by the EU FP7 Marie Curie Zukunftskolleg Incoming Fellowship Programme, University of Konstanz (grant no. 291784). The Swiss National Science Foundation has supported Carl Müller-Crepon and Yannick Pengl through grants P0EZP1\_165233 and P0EZP1\_159076, respectively.

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