# Coethnics Covote in Africa: Studying Electoral Cleavages with a Covoting Regression Model

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Carl Müller-Crepon* Nils-Christian Bormann†

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**Abstract**

Ethnicity is an important cleavage in Africa, yet its influence on voting is contested. Selection biases from restricted choice sets complicate micro-level analyses, while bias from ecological inferences and unobserved confounders hamper meso and macro-level approaches. Our new *Covoting Regression* (CVR) tackles several of these challenges. It estimates the effect of coethnicity on the probability that pairs of voters covote for the same party/candidate while conditioning on other characteristics shared between them. Thereby, CVR mirrors the micro-foundations of widely-used aggregate indicators, such as the Herfindahl-Hirschman indices of party concentration and ethnic homogeneity. Our data consists of dyadic comparisons between respondents from Afrobarometer surveys. Pooling across 28 countries, coethnicity increases covoting intentions by 17 percentage points. The effect of coethnicity is partially driven by politically relevant groups and covoting for ethnic parties. It is consistent across institutionally diverse countries and at least five times larger than that of other cleavages. Beyond ethnicity, we address key issues in studying electoral consequences of socio-economic cleavages and bridge gaps between levels of analysis.

**Keywords:** Cleavages; Voting; Ethnic voting; Ethnic parties; Ethnicity; African politics; Computational Methods

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*Department of Government, London School of Economics and Political Science. Corresponding author: c.a.muller-crepon@lse.ac.uk

†Department of Philosophy, Politics & Economics, Witten/Herdecke University. nils.bormann@uni-wh.de

It used to be a truism among political scientists that African voters support coethnic candidates and African parties target coethnics in election campaigns (e.g., Horowitz 1985; Rabushka and Shepsle 1972; Bayart 2009).[1] Increasingly, however, researchers question this hypothesis. Some micro-level studies propose a broader sociological understanding of voting by identifying alternative cleavages like religion or urban-rural differences (e.g., McCauley 2014; Nathan 2016); others adopt rationalist perspectives that emphasize individual economic interests and the quality of information available to voters (e.g., Bratton, Bhavnani and Chen 2012; Casey 2015; Ferree, Gibson and Long 2021). Meso and macro-level analyses identify electoral systems and ethnic inequality as conditioning factors of ethnic voting (Huber 2012; Huber and Suryanarayan 2016). Some studies even suggest that the effects of ethnicity on vote choice and party systems in Africa are in part spurious, simply reflecting underlying geographic clustering or shared economic interests (Ferree and Horowitz 2010; Boone et al. 2022).

Yet, micro, meso, and macro-level analyses of the electoral effects of ethnic cleavages suffer from potentially severe methodological problems. At the micro-level, the interdependence between observed voting patterns and the fixed menu of parties and candidates in any one election complicates inference. If researchers conceptualize ethnic voting as the support of voters for coethnic candidates, they risk *selection bias* if some voters exclusively face (non-)coethnic candidates (see also Ferree 2022). Moreover, the idiosyncratic and ever-changing menu of parties or candidates in individual elections prevent comparisons across countries and elections. Meso and macro-level comparative approaches address this challenge by analyzing the degree to which ethnic groups differ in their voting patterns (Huber 2012; Huber and Suryanarayan 2016; Houle, Park and Kenny 2019). Yet, inferring individual-level voting motivations from group or country-level analyses constitutes a clear case of *ecological inference*. Since ethnic cleavages frequently correlate with other social divisions such as geographic or economic differences, this approach might be biased by omitting non-ethnic determinants of vote choice (Ferree and Horowitz 2010; Boone et al. 2022).

[1] Our argument focuses on elections of parties and candidates, which is why we use these labels interchangeably.

We introduce a Covoting Regression (CVR) model as a new analytical approach. The CVR combines the complementary strengths of micro-level voting studies and comparative work at the meso and macro-levels to address several of the problems affecting either. Similar to aggregate indices employed by comparativists, the model assesses the likelihood of covoting intentions among pairs of individuals, which we define as the shared intention to vote for the same party or candidate.[2] Yet, instead of building aggregate country or group-level measures, we model covoting at the micro-level of individual *pairs of voters*.[3] This allows us to estimate the effect of coethnicity (or any other cleavage) between individuals conditional on other dyadic socio-economic characteristics with a standard linear probability model.

Conceptually, the estimated conditional effect of coethnicity combines Huber's (2012) group- and party-based understanding of ethnic voting. It increases with greater differences in voting patterns between ethnic groups *and* greater differences in the ethnic composition of parties' voters. Aggregating across these two dimensions, the CVR thus explicitly considers ethnic bloc voting for candidates from other ethnic groups as an integral part of ethnic voting. Conceptually, this is a major departure from prevalent micro-level approaches, which typically rely on coethnicity between voters and candidates. Beyond the broad micro-level interpretation, we use the CVR to derive covariate-adjusted estimates of group-based and party-based conceptualizations of ethnic voting at the meso-level (Huber 2012). At the macro-level, the model's estimates directly reflect the classic Herfindahl-Hirschman Index and can be interpreted as the elasticity of party system concentration to marginal changes in ethnic homogeneity. The CVR thus bridges the gap between the differing levels of analysis in studies of electoral behavior with possible applications to cleavages beyond ethnicity.

The CVR solves several of the aforementioned methodological problems of micro, meso, and macro-level approaches. By modeling pairwise covoting, parties and candidates disappear from our formulation. This abstraction reduces selec-

[2]While we empirically measure covoting intentions and party preferences, our model applies conceptually to actual covoting. We use the shorter term in our theoretical discussion.

[3]Our dyadic approach is similar to work in international relations on the formation of trade blocs or military alliances.

tion biases from restricted choice sets and allows for comparative analyses across countries and over time. It does, however, come at the cost of the model's inability to yield group- or party-specific inferences. In addition, other forms of selection bias, resulting from missing data and endogenous party formation, still affect the CVR.[4] At the same time, estimating the conditional association between coethnicity and covoting addresses ecological inference problems and bias from confounders which can affect meso- and macro-level studies, while increasing statistical power.

Empirically, we use the CVR to provide comprehensive evidence on the effect of coethnicity on covoting across Sub-Sahara Africa. Drawing on multiple rounds of the Afrobarometer surveys from 28 states across Sub-Sahara Africa, we recast each country sample into pairwise comparisons between respondents. We proxy unobserved actual covoting by measuring our *outcome variables* as covoting intentions in presidential elections and shared preferences for a political party.[5] Coethnicity, the main *explanatory variable*, is measured as the shared mother tongues of respondents. Shared demographic, economic, and geographic characteristics as well as survey-round-fixed effects constitute our controls. Building on our conceptual derivation of the CVR, we estimate the probability of covoting intentions in linear probability models, and adjust standard errors to address the repeated inclusion of individuals from a limited number of ethnic groups across dyads.

Our results show strong support for the dominance of coethnicity in shaping covoting intentions in most African states in our sample. We estimate that coethnicity by mother tongue increases the probability that two respondents have shared voting intentions by 17 percentage points. Zooming in on specific countries and elections, we discuss variation in the estimated effect of coethnicity on covoting intentions over time and across cases. Additional analyses account for cross-cleavage interactions, alternative measures of coethnicity and linguistic distance, different sampling procedures, and modeling choices. We even uncover substantial, if smaller, effects of coethnicity on covoting intentions within comparatively homogeneous villages or neighborhoods, suggesting that coethnicity captures more than shared, spatially clustered political preferences.

[4]Note however that the reflection of ethnic bloc voting for non-coethnic candidates alleviates biases from endogenous party formation in some but not all contexts.

[5]The Afrobarometer asks survey participants which party they feel close to.

In a set of descriptive analyses, we demonstrate how the CVR can be applied to test explanations of ethnic voting or changes therein at the three levels of analysis. At the micro-level, we find that higher ethno-centric trust and shared perceptions do not robustly explain coethnic covoting. At the meso-level, the effect of coethnicity is augmented in politically relevant ethnic groups and strongest for covoting for ethnic parties. At the macro-level, we find no substantive moderating effects of electoral, democratic, or traditional institutions.

Beyond coethnicity, our results also show positive effects of religious, educational, occupational, and geographic similarities on covoting, thus supporting conjectures on socio-economic characteristics (e.g., Boone et al. 2022; Bratton, Bhavnani and Chen 2012; Koter 2016; McCauley 2014). Yet, across our broad sample of African electorates, the effects of these alternative cleavages are at least five times smaller than that of coethnicity.

We conclude by discussing the implications of our new approach for the study of electoral cleavages. In doing so, we highlight the importance of bridging levels of analysis to achieve inferences of high internal and external validity and discuss the utility of the CVR in studying the electoral effect of socio-economic cleavage. Potential extensions can apply the CVR to local election results and individual attitudes and choices beyond voting.

## Ethnicity and Voting in Sub-Sahara Africa

Political scientists have come a long way from the once paradigmatic view that elections in Africa constituted an ethnic census (Horowitz 1985, 196). Classic works on vote choice in Sub-Saharan Africa either stress psychological or instrumental motivations for ethnic voting and the corresponding existence of ethnic parties (Mozaffar, Scarritt and Galaich 2003). In short, the psychological approach entails that voters reaffirm their identity through voting for coethnic candidates and attempt to avoid discrimination by ethnically distinct rulers. Political leaders cannot escape the logic of ethnic outbidding, in which more extreme political demands on behalf of coethnics increase electoral support (Rabushka and Shepsle 1972; Horowitz

1985).[6] In turn, instrumentalists suggest that African voters support coethnic candidates to receive economic benefits through clientelist exchanges during the election period and patronage distribution afterwards, if "their" candidate joins the ruling coalition. Political elites themselves prefer to build ethnically-based support coalitions to limit access to state funds to ethnic insiders (Bates 1974).

Political scientists investigate individual-level demand and party-level supply at the micro, meso, and macro-levels. *Micro-level* studies probe both the mechanisms that link individuals' ethnic identity to vote choice and alternative explanations for party preferences. Experimental research supports an instrumentalist interpretation of ethnic voting that portrays the ethnicity of candidates as an informational shortcut for voters to learn about the likelihood of future economic benefits (Ferree 2006; Ferree, Gibson and Long 2021; Conroy-Krutz 2013; Carlson 2015; Casey 2015; Ferree, Gibson and Long 2021).[7] In contrast, survey-based research frequently explores alternative voter motivations, such as economic performance evaluations and education (Bratton and Kimenyi 2008; Bratton, Bhavnani and Chen 2012), partisanship (Ferree and Horowitz 2010; Hoffman and Long 2013), populism (Resnick 2012), and urban-rural differences (Nathan 2016, 2019). Others stress that ethnic voting depends on local factors characteristic for many African countries, including traditional authorities (Baldwin 2013; Koter 2016), and local ethnic geography (Ichino and Nathan 2013).

Comparativists working at the *macro-level* complement micro-level studies by highlighting supply factors. Guided by Duverger's law, cross-country comparisons suggest that the effective number of ethnic groups in a country correlates with the effective number of parties under permissive electoral rules (Mozaffar, Scarritt and Galaich 2003; Clark and Golder 2006; Lublin 2017). In contrast, case studies show that party competition goes beyond ethnicity (Elischer 2013) and that presidential candidates pursue non-coethnic supporters (Horowitz 2022). Larger groups are indeed more likely to have coethnic candidates on the ballot (cf. Posner 2004*b*) and to see intra-group competition Ferree (2010). Echoing micro-level insights on the im-

[6]Sociological explanations of voting following Lazarsfeld, Berelson and Gaudet (1968/1944) are curiously absent from the ethnic voting literature, a theme we return to in the conclusion.

[7]Though see Adida et al. (2017), who highlight motivated reasoning as linking ethnic identity and performance evaluation. Alternatively, Mor (2022) documents grievance-driven demands for a Catholic party in 19th-century Prussia.

portance of alternative cleavages, Boone (2024) argues that the correlation between ethnicity and vote choice hides the effects of regionally concentrated economic interests (see also Ishiyama 2012). Analyses of ethnic voting at the group or *meso-level* falls in between these two end points by highlighting group-level economic inequality and electoral rules as moderators (Huber 2012; Huber and Suryanarayan 2016; Houle, Park and Kenny 2019).

Studies at each level of analysis complement each other. Substantively, research at all levels raises questions about the influence of other cleavages including urban-rural differences (Nathan 2016; Wahman and Boone 2018; Harding 2020), regional coalitions (Ferree and Horowitz 2010; Boone 2024), and economic class interests (Bratton, Bhavnani and Chen 2012; Resnick 2012; Huber and Suryanarayan 2016; Houle, Park and Kenny 2019). Micro-level research trades in generalization against detailed, case-specific insights with the potential of identifying causal underpinnings of vote choice. Meso and macro-level comparisons potentially offer greater external validity but rely on *ecological inference* and risk omitted variable bias.[8] Taking groups as units of analysis or building blocks of aggregate indices, meso and macro-level analyses assume the primacy of ethnicity rather than testing it.[9]

Importantly, micro, meso, and macro-level approaches differ in their conceptualization of ethnic voting itself. Both macro and micro-level studies commonly analyze the degree of ethnic voting as the probability that an individual chooses a coethnic candidate or party. Meso-level research, in turn, analyzes the extent of ethnic divergence in voting patterns based on the electoral distance between groups and ethnic distance between parties (Huber 2012). Respectively, we label these alternative strategies the *voter-candidate correspondence* and the *divergence* approach. In the correspondence approach, ethnic voting increases as members of a group increasingly vote for parties that represent their group *and* their proportion among all voters of these parties increases. In contrast, ethno-political divergence increases under two alternative conditions: as a political party increasingly draws support from one ethnic group in what Huber (2012) coins the "party-based approach", *or*

[8] Appendix B.2 illustrates this bias via simulations.

[9] Fixing group boundaries also precludes a more nuanced understanding of potentially variable ethnic boundary markers or even endogenous processes of identity change (Posner 2005; McCauley 2014; Green 2021; Müller-Crepon 2023).

as voters from one ethnic group vote together ("group-based approach").

This contrast brings into relief three important conceptual and methodological issues. First, empirical analyses that employ the correspondence approach prioritize a party- rather than group-based understanding of ethnic voting. This can lead to an underestimation of ethnic voting when members of a group vote en bloc for a non-coethnic candidate. While the correspondance approach would not regard those votes as "ethnic", the divergence approach would detect this ethnicization of political preferences.[10] Second, the correspondence approach classifies votes for coethnics as ethnic even if they were cast for a "non-ethnic" reason or even due to pure randomness. This produces an upward bias in aggregate estimates of ethnic voting.[11] Third, constrained choice sets can leave voters with either exclusively coethnic or non-coethnic candidates (Ferree 2022). This produces selection bias in the correspondence approach.[12]

This selection bias arises in three analytical situations. First, potential candidates do not run in expectation of low support from their group, for example, where groups are too small to lift candidates to victory (Posner 2004*b*). When excluding such groups from an analysis, analysts will potentially overestimate ethnic voting in a society. Second, the utility of ethnic voting might be stable across groups, but some groups might still have no coethnic candidate. Here, a "0" classification of ethnic voter-candidate correspondence mixes the absence and the impossibility of ethnic voting. If members of such "unrepresented" groups voted en bloc for a "second-best" non-coethnic who represents their interests, correspondence approaches would underestimate the overall extent of ethnic voting. Huber's divergence approach, however, can capture group-based voting for non-coethnic candidates. In a third situation, a broad choice set is available to voters but the analyst artificially limits vote choice, for example, by only evaluating support for the incumbent (e.g., Bratton and Kimenyi 2008; Bratton, Bhavnani and Chen 2012). In this case, the degree of ethnic voting may be over- or underestimated depending on the utility that individuals derive from ethnic voting and the distribution of

[10] Appendix A formalizes the differences between the various approaches.

[11] See Appendix A.3 for a simulation of this bias.

[12] Selection bias affect experimental studies less, but biases may occur if they are inconsistent with the real-world experiences of participants (see Dafoe, Zhang and Caughey 2018).

votes for excluded candidates.[13]

Because we only observe real but not hypothetical alternative candidate choice sets, those analyzing observational data currently have two options.[14] First, voting studies that rely on the correspondence approach would need to adjust for factors that influence the existence of "ethnic" candidates (Ferree 2022), and candidates' expected vote shares under non-ethnic voting. However, this strategy is difficult to implement in broadly comparative work. Second, the divergence approach developed by Huber (2012) offers the possibility to model group-based voting across contexts even where choice sets are constrained. It frees the researcher from specifying correspondences between voters and parties, and from taking arbitrary decisions on which groups to include in the analysis. However, the aggregation of individual choices to the group level makes this approach vulnerable to omitted alternative cleavages.

In sum, a new approach to observational research on ethnic voting should (1) account for alternative motivations and identity categories that might serve as a basis for voting to avoid omitted variable bias; (2) avoid the pitfall of artificially limited choice sets to overcome selection bias, and (3) facilitate coverage of a broad number of countries over time to ensure external validity.

## Covoting Regression: A new way to estimate the effects of (ethnic) cleavages

Our new Covoting Regression (CVR) bridges prevailing approaches to ethnic voting by combining their strengths which together address their respective weaknesses. In short, we follow macro- and meso-level measurement strategies to conceptualize the effect of ethnic cleavages as the effect of coethnicity on covoting intentions. Overcoming reliance on aggregate indicators which leads to problems of ecological inference and omitted variable bias, we study covoting intentions among pairs of individuals contained in survey data with standard regression models. In

[13] Appendix A.3 simulates scenarios 1 and 2. We show that the correspondence approach entails much larger selection bias than our new covoting approach.

[14] Experimental research is in turn better placed to address the problem of endogenous candidate supply which is typically controlled by the researcher.

the following, we introduce our method, explain the transformation of the Afrobarometer data into voter pairs, and discuss the empirical CVR specification.

### Bridging micro, meso, and macro-approaches

Macro-approaches often study the structure of party systems as the result of an interplay between institutional determinants and societal cleavages. Empirically, they employ aggregate measures to operationalize the main variables of interest, such as the Herfindahl-Hirschman Index (HHI) and the effective number of parties (ENP) to measure party system concentration, or the HHI of ethnic homogeneity. Yet, while the empirical focus is on the macro-level, these measures have explicit meso- and micro-foundations in theory and measurement. In particular, the HHI of party concentration and its inverse, Laakso and Taagepera's (1979) ENP are often computed with meso-level data on parties' vote shares as

$$HHI(party) = ENP^{-1} = \sum_{p=1}^{K} \left( \frac{N_p}{N} \right)^2, \tag{1}$$

where $\frac{N_p}{N}$ is the vote share of party $p \in K$ among voters $N$. Party concentration thus decreases with more and more equally sized parties, which in turn increases the effective number of parties. As scholars of ethnic fragmentation note, the HHI has clear micro-foundations: it reflects the chance that two randomly drawn individuals belong to the same group or category (e.g., Alesina, Baqir and Easterly 1999). In terms of voting, the relative vote shares of one party simply reflect individuals' voting probability for that party. The square of that probability then yields the chance that two random voters voted for the same party. We can thus reformulate $HHI(party)$ as the chance that two randomly chosen voters vote for the same party:

$$HHI(party) = \frac{1}{N^2} \sum_{i,j}^{N} \mathsf{covoting}_{i,j}, \tag{2}$$

where $i$ and $j$ are individual voters drawn from all voters $N$, and $\mathsf{covoting}_{i,j}$ is an indicator that returns 1 if $i$ and $j$ covote for the same party $p$ and 0 otherwise.

Rarely discussed, the covoting formulation in Equation (2) includes comparisons within the same individual $i = j$, inducing downward bias when computed on a finite sample of individuals $N$. This is because comparisons within the same voter $i = j$ must yield $\textsf{covoting}_{i,j} = 1$.[15] Drawing on Simpson (1949), this bias can be corrected by avoiding "within-individual" comparisons when computing the HHI:

$$HHI(party) = \sum_{p=1}^{N}(\frac{N_p}{N}\frac{N_p - 1}{N - 1}) \tag{3}$$

$$= \frac{1}{N^2 - N}\sum_{i,j,i\neq j}^{N} \textsf{covoting}_{i,j} \tag{4}$$

Equation 4 shows how the unbiased expectation of $HHI(party)$ is equivalent to the average *covoting* rate between all voters.[16] In turn, ethnic homogeneity among voters as measured through the HHI is equivalent to the average rate of coethnicity between all voter pairs. Moving beyond measures of diversity, measures of dispersion such as the Gini coefficient can be similarly reformulated as comparisons between pairs of individuals.[17]

We hone in on these micro-foundations of macro-level approaches and propose to model the effect of ethnic cleavages on voting by estimating the effect of coethnicity on covoting in pairs of individuals $i$ and $j$. We start deriving our *Covoting Regression* (CVR) model as

$$\textsf{Covoting}_{i,j} = \beta_0 + \epsilon_{i,j} \tag{5}$$

with $i, j \in N, i > j$.[18] In this formulation, estimate $\beta_0$ captures the average rate of covoting among all pairs of individuals and is thus equivalent to $ENP^{-1}$ in Equation (4).[19]

The pairwise regression model in Equation (5) can be easily extended by adding a covariate matrix $\mathbf{x}_{i,j}$ which measure individuals' $i$ and $j$'s similarities or differ-

[15]The bias decreases with $N$.

[16]Equation (4) approaches Equation (2) as $N$ increases towards infinity.

[17]The Gini coefficient can be computed as half the mean absolute (wealth, income, education, etc.) difference among all pairs of individuals, see e.g., Sen (1997, p. 31).

[18]As covoting is not directional, we limit ourselves to one comparison between any two voters. Computationally though, the result is the same as in Equation 4.

[19]The estimate of the intercept in an empty linear regression model, $\beta_0$, equals $\bar{y} = \frac{1}{n}\sum_i^n y_i$ (Wooldridge 2008, 29). Appendix Figure A7 empirically demonstrates this equivalence. Our estimation method of ENP and HHI adds the benefit of uncertainty estimates reflecting survey sampling.

ences on important socio-economic cleavage dimensions as well as other control variables such as cross-cleavage interactions.[20] We thus propose to estimate the effect of coethnicity on covoting as

$$\text{Covoting}_{i,j} = \beta_0 + \beta_1\, \text{coethnicity}_{i,j} + \gamma\, \mathbf{x}_{i,j} + \epsilon_{i,j}. \tag{6}$$

The focus of the CVR on covoting between $i$ and $j$ abstracts from the identities of candidates and parties. Their disappearance from the formulation circumvents the selection bias that arises in the correspondence approach when linking candidates to ethnic groups. Yet, the CVR remains vulnerable to biases from endogenous candidate supply when analyzing observational data. However, we find that its accounting for ethnic bloc voting decreases the size of this bias in comparison to the correspondence approach.[21] We also note that the CVR can be used to evaluate data from experimental studies with exogenous candidate supply.

Due to the equivalence of the $\text{coethnicity}_{i,j}$ indicator in Equation (6) with the micro-level interpretation of the HHI index, the estimate for $\beta_1$ has interpretations at all levels of analysis. At the macro-level and conditional on covariates $\mathbf{x}_{i,j}$, it is the elasticity of the party concentration in response to marginal changes in ethnic homogeneity[22] such that

$$\frac{\delta\, HHI(party)}{\delta\, HHI(ethnic)} = \frac{\delta\, \text{covoting}_{i,j}}{\delta\, \text{coethnicity}_{i,j}} = \beta_1, \tag{7}$$

This feature of the CVR extends to other dyadic comparisons between voters that reflect macro-level measures, such as pairwise wealth differences which together constitute the Gini coefficient.

At the meso-level of pairs of ethnic groups, $\beta_1$ captures the difference in average covoting within groups and covoting between groups, with group-pairs weighted by the product of their size. As a consequence, Eq. 6 can also be used to derive the expected rate of covoting among coethnics as well as the expected rate of coethnic-

[20]Appendix B.2 shows how the CVR successfully accounts for correlated cleavages.

[21]See Appendix A.3.

[22]Interpretations that extrapolate beyond small margins neglect the possibility of corresponding strategic party and exit and its knock-on effects on covoting among (non-)coethnics. See Appendix B.1.

ity among covoters at average covariate values $\bar{\mathbf{x}}$:[23]

$$Pr(\text{covoting} \,|\, \text{coethnicity} = 1) = \beta_0 + \beta_1 + \gamma\bar{\mathbf{x}} \tag{8}$$

$$Pr(\text{coethnicity} \,|\, \text{covoting} = 1) = \frac{\overline{\text{coethnicity}}}{\overline{\text{covoting}}}\,(\beta_0 + \beta_1 + \gamma\bar{\mathbf{x}}), \tag{9}$$

where the rate of coethnicity among covoters in Eq. 9 is the rate of covoting among coethnics scaled by the relative share of coethnic pairs compared to covoters in the population. These measures closely resemble the inverse of Huber's (2012) measures of ethnic Group Vote Fractionalization and Party Vote Fractionalization (see Appendix D.2). Importantly, the conditional rates of coethnicity and covoting are driven not only by the effect of coethnicity on covoting $\beta_1$, but also by the intercept as well as average covariate values and effects.

Finally, at the micro-level, $\beta_1$ can be interpreted as the marginal effect of coethnicity on the probability of covoting between individuals, again conditional on covariates. In principle, this covoting probability can also be recovered from voting probabilities derived from a multinomial logistic or linear regression model (Appendix A.4). Doing so yields largely equivalent results but does not yield standard uncertainty estimates and comes with estimation problems where data is sparse. Reformulating covoting as the product of two individual probabilities to vote for the same party further reveals how to address biases that stem from the endogeneity of one cleavage $x_1$ to another $x_2$. Where such endogeneity exists, it becomes necessary to not only model voter characteristics of $i$ and $j$, but also their interactions $x_{1,i}\, x_{2,j}$ and $x_{2,i}\, x_{1,j}$, thereby capturing cross-cleavage covoting.

By providing interpretations at all levels of analysis based on dyadic, micro-level modelling, the CVR combines the advantages of previous approaches. In particular, it avoids making ecological inferences from macro-level modelling but at the same time allows for inter-temporal and cross-country comparisons of the conditional association of coethnicity with covoting. While the conceptualization of ethnic voting as coethnic covoting entails some loss of detail, it prevents biases from narrower but potentially more informative definitions of ethnic voting which

[23]See Appendix A.2 for further discussion and parallel derivations of the rate of covoting among non-coethnics and non-covoting among coethnics which have comparative value.

require ex ante coding or standardization of parties or candidates. We show in our analysis of heterogeneous effects and mechanisms how the CVR can be used to test theoretical arguments about the drivers of ethnic voting at all three levels of analysis. To facilitate understanding of our main empirical strategy, we first introduce our data structure and then present our regression model.

### Dyadic data on covoting, coethnicity, and other cleavages

To operationalize our analysis of covoting intentions among individuals, we transform survey data into pairs of individual respondents. For each pair, we encode whether respondents share voting intentions and measure ethnic and other cleavages through respondents' pairwise shared ethnicity and similarity in other socioeconomic characteristics.

Our main data source consists in the nationally and, in expectation, locally representative Afrobarometer survey series which contains data on political preferences across an increasingly large set of states in Sub-Sahara Africa since 1999. For the most part, we rely on the survey's seventh round fielded between 2015 and 2018 in 28 states with 1'200 respondents per country.[24] To gauge variation in the effect of ethnic cleavages over time, we draw on rounds 3–7.[25] In addition to surveying preferences for presidential candidates and political parties, the surveys cover a large range of demographic and economic items, and provide geographic information on respondents' place of residence. The resulting information allows us to capture covoting intentions along with a wide range of non-ethnic cleavage dimensions discussed by existing work.

**Unit of analysis:** Following the logic introduced in Equation (4), we transform survey-rounds for each country into undirected dyadic comparisons between all respondents $i, j \in N_{c,t}$ with $i \neq j$. This gives rise to a total of $(N_{c,t}(N_{c,t} - 1))/2$

[24] Benin, Botswana, Burkina Faso, Burundi, Cameroon, Côte d'Ivoire, Ethiopia, Gabon, Gambia, Ghana, Guinea, Kenya, Lesotho, Liberia, Madagascar, Malawi, Mali, Mozambique, Namibia, Niger, Nigeria, Senegal, Sierra Leone, South Africa, Tanzania, Togo, Uganda, Zambia, Zimbabwe. We randomly subset the larger surveys from Ghana, Kenya, Mozambique, Nigeria, and South Africa to 1'200 to give each country roughly equal weight.

[25] Rounds 1 and 2 do not include an item on preferences over candidates in potential presidential elections.

observations per country-round. After dropping observations with missing data, our main analysis of preferences for presidential candidates (parties) draws on a median number of 733 (536) respondents and 268′278 (143′380) dyadic comparisons between them per country surveyed.[26]

**Measuring covoting intentions:** Because actual voting behavior remains unobserved in the Afrobarometer, we encode our main two measures of covoting intentions by drawing on answers to questions on respondents' preferences over presidential candidates and parties:

> Voting intention: If a presidential election were held tomorrow, which party's candidate would you vote for?
>
> Party preference: Do you feel close to any particular political party? Which party is that?[27]

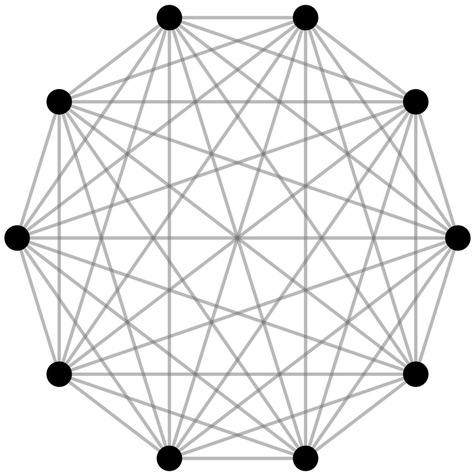
Drawing on these items, we record two dummy variables that take the value of 1 if respondents $i$ and $j$ share a preference for the same candidate or party and 0 otherwise.[28] The result is visualized for a sample of 10 respondents from Ghana in Figure 1. Each dot represents one respondent with its color reflecting their preferred candidate/party. Lines between respondents are drawn in black ('1') where they share a preference and in grey ('0') where they do not. On average, we observe that respondents share their voting intention and party preference in 45.9 and 45.5 percent of all dyads, which corresponds directly to the cross-country average HHI of party concentration or the inverse ENP (see Equation 4).

**Coethnicity:** We capture our main explanatory variable of interest – respondents' pairwise coethnicity – in a binary variable that records whether they share the same

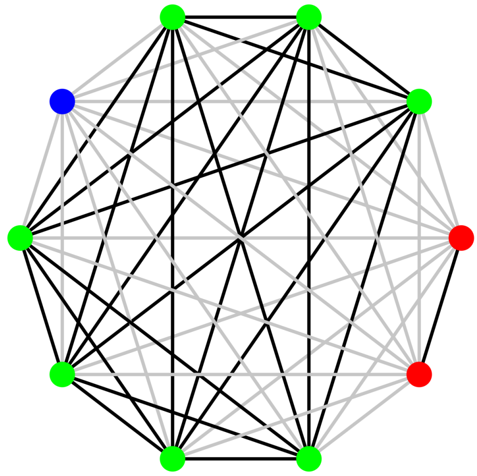
[26] Appendix Figure A18 shows that our results are robust to reducing the number of comparisons per respondent down to as few as one. Weighing observations such that each country-round receives equal weights slightly increases our main estimates, see Appendix Table A19.

[27] Missing values are recorded for respondents who do not feel close to any party.

[28] We drop individuals with missing responses and "other" parties as missing. While sensitive party identification in polarized environments can lead to selection bias (Davis and Wilfahrt 2025), results are robust to coding "other" parties as separate parties for each respondent (Appendix Table A20).



(a) Fully connected graph



(b) Shared voting intentions

Figure 1: Covoting dyads from 10 respondents in Ghana, Round 7

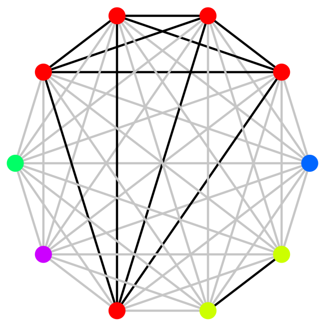
Note: In (b): NPP green, NDC red, PPP blue. Covoting as black edges.

mother tongue (1) or not (0).[29] As visualized for the randomly drawn 10 Ghanaians in Figure 2a, this leads to many coethnic dyads among respondents from large language groups (e.g., the Akan in red). On average, respondent pairs share their mother-tongue at a rate of 28 percent, which corresponds to the average HHI of ethnic homogeneity across the countries in the sample.

Respondents' reported mother-tongue is among the least malleable ethnic identity indicators and therefore least likely affected by reverse causality or omitted variable bias. In contrast to the language spoken at respondents' home or their self-proclaimed ethnic identity, mother tongues are not affected by political concerns of respondents (e.g. Green 2021) and assimilation over their life-course (e.g. Müller-Crepon 2023).[30] We employ three different strategies to address the remaining potential for omitted variable bias through, for example, economic factors affecting political preferences as well as ethnic identities, and reverse causation, such as multi-generational assimilation that aligns ethnic to political identities. First, we condition our estimates on several other individual-level covariates which might

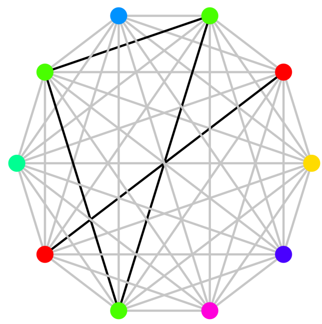
[29]The respective question reads: "Which [enter nationality] language is your mother tongue or language of origin?" Note that Afrobarometer round 7 is the first to ask specifically about respondents' mother tongue as separate from the language spoken in their home *now*. Hampering comparisons over time, all previous rounds ask about respondents' "home language" which leaves this crucial distinction open.

[30]We test the alternative measurements of ethnic identity in Appendix F.3.



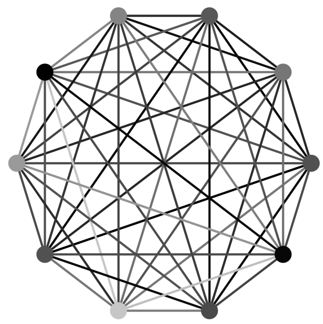
(a) Shared mother-tongues

Note: Colors denote mother-tongues, intra-group edges in black



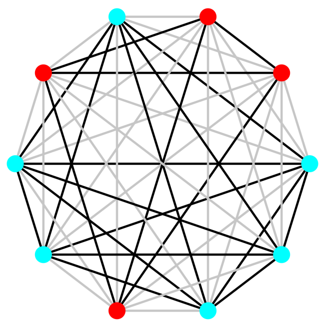
(b) Shared religion

Note: Colors denote religious groups, intra-religion edges in black



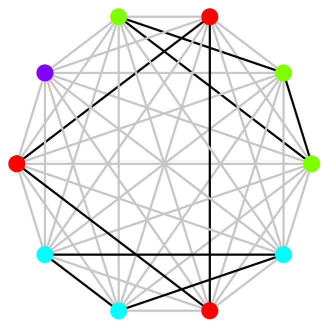
(c) Age similarity (decades)

Note: Grey-scale denotes older age and greater similarity



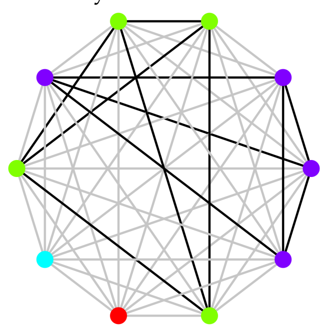
(d) Shared gender

Note: Female in light blue, male in red, intra-gender edges in black



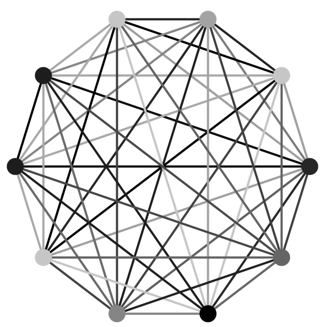
(e) Shared education

Note: Colors denote education levels, intra-education level edges in black



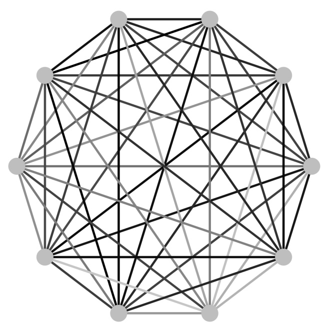
(f) Shared occupation

Note: Colors denote occupations, intra-occupation edges in black



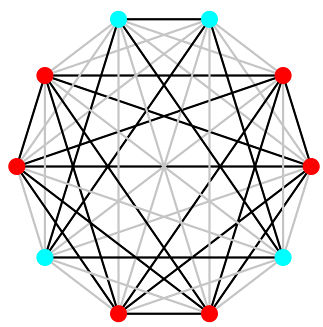
(g) Wealth similarity

Note: Dark shades denote wealthier respondents and greater similarity



(h) Geographic proximity

Note: Darker shades denote greater proximity



(i) Shared urban-rural status

Note: Urban blue, rural red, shared status edges in black

Figure 2: Encoding of main explanatory variables on example graph of 10 respondents from Ghana

affect individuals' stated ethnic origin and their political preferences. Second, we analyze variation in effects at short and large linguistic distances, which are harder to overcome through assimilation or misrepresentation. Third, a set of robustness checks zooms in on covoting intentions among respondents from the same enumeration area, thus holding geographic factors constant.

**Control variables:** With regard to our first strategy of conditioning on observables, we encode a set of pairwise comparisons between respondents that capture prominent political cleavages and might affect individuals' reported language. All are visualized for our exemplary 10 Ghanaian respondents in Figures 2b to 2i. For reasons of consistency, we construct our measures such that larger positive values denote greater similarity between respondents which should, in expectation, come with higher probabilities of covoting intentions.

First, we complement our measure of shared mother tongues by accounting for whether respondents share the same *religion*. Second, we capture *demographic similarities* between respondents by recording age and gender similarities. Third, we approximate *economic cleavages* by adding dummy variables for shared educational and occupational background as well as wealth similarity, measured as one minus absolute wealth differences.[31] Lastly, we capture purely *geographic cleavages* by including as-the-crow-flies proximity between respondents (in 1'000km) and a dummy variable capturing whether respondents share their urban or rural status.

**Combining data across countries and rounds:** Because our measures of covoting intentions, coethnicity, and additional control are measured as binary or continuous indicators of similarity the data can be stacked and analyzed across countries and rounds without any additional processing. This is a substantive advantage over standard approaches of modeling the effect of ethnic (or other) cleavages on party or candidate preferences which require harmonization across context with the selection biases this gives rise to.

[31] We compute wealth differences from an individual-level wealth index constructed with a principal component analysis of respondents' availability of food, water, healthcare, and income. As noted above, the resulting measure is closely related to the Gini coefficient. An extended version of the index adds respondents' level of education, employment type, and availability of fuel and yields very similar results (Appendix F.5).

### Modelling the effect of ethnic cleavages on covoting

With the undirected dyad of respondents $i$ and $j$ as our main unit of analysis, we employ a linear probability model to estimate the probability of covoting intentions between respondents along the lines of the CVR proposed in Equation 6.[32] We move beyond the confines of one survey sample of individuals and generalize the model across countries and time as

$$\text{Covoting}_{i,j,c,t} = \alpha_{c,t} + \beta_1 \, \text{coethnic}_{i,j} + \gamma \, \mathbf{x}_{i,j} + \epsilon_{i,j} \tag{10}$$

where dyads are constructed only among individuals observed in the same country $c$ at time $t$. $\alpha_{c,t}$ is a fixed effect for each country $c$ at time $t$, capturing the average rate of covoting in a country and the effects of unobserved country-level factors, such as states' population size, history, or electoral institutions. $\beta_1$ then captures the *average* effect of coethnicity by mother tongue on covoting intentions among individuals across all countries and times in the sample, conditional on the covariates **x** introduced and visualized in Figure 2 above.

Since all socio-economic factors underlying $\mathbf{x}_{i,j}$ can plausibly be causes *and* consequences of respondents' ethnic identity, adding these controls could trigger post-treatment bias. At the risk of omitted variable bias, we therefore first estimate a baseline model without any controls. We then compare the estimates across the baseline and the model with controls. Although it is theoretically possible that the potential biases from omitted variables and post-treatment controls lead us to over-estimate the effect of coethnicity in both cases, small differences between the two specifications and subsequent robustness checks would suggest that we approximate the true effect of coethnicity.

By construction, the CVR model is estimated on interdependent data. We thus consider various strategies to adjust standard errors. In our main analyses, we rely on the conservative two-way clustering on the ethnicity of individuals $i$ and $j$ that constitute each dyad. These clusters correspond to the level of "treatment assignment" if we consider ethnic groups and their members to be jointly treated as groups. We explore alternative clustering strategies in Appendix F.7: Clustering

[32] Logistic regression models yield equivalent results. See Table A21.

at the country-level yields just as conservative standard errors, while clustering on the level of individuals or their enumeration area yields smaller error bands. Lastly, we show that Aronow et al.'s (2015) cluster-robust variance estimator for dyadic data yields comparatively small uncertainty estimates when employed at the level of individuals, their ethnicity, or their locations of residence.[33]

## Results

Our analysis yields strong support of the hypothesis that coethnicity increases the rate of shared voting intentions and party preferences among among Afrobarometer respondents. Table 1 presents our main estimates, showing the unconditional and conditional effect of coethnicity on respondents' covoting intentions, measured as shared voting intentions and preferences for parties. We find that pairs of respondents who share their mother tongue are between 16.4 and 17.8 percentage points more likely to have the same voting intention and party preference. The effect is very stable across specifications, does not vary between our two outcomes, and is associated with little uncertainty ($p < .001$).

Substantively, these effects are large. We observe shared covoting intentions in 46 percent of all survey respondent dyads. The conditional increase in covoting intentions resulting from shared mother-tongues of 17 percentage points (Models 2 and 4) thus amounts to 37 percent of the average rate of covoting. The estimated effect of coethnicity also swamps that of any other pairwise similarity measure between respondents, the substantively largest being that of shared occupation with an estimated effect of 2.8 percentage points. We compare the effect of coethnicity with other cleavage dimensions more thoroughly below.

At the meso-level, the coefficient of coethnicity together with the base rates of covoting intentions and coethnicity allow for computing the group- and party-based measures of ethnic voting (Eqs. 8 and 9). Based on Model 4, 57.6 percent of coethnics covote compared to 41.2 of non-coethnics. In turn, adjusted for covariates we estimate that there are 34.7 percent of coethnics among covoters, compared

[33]Beyond its effects on uncertainty estimates, unit-interdependence may bias point estimates in our setting. When explicitly addressing this concern with a *Probabilistic Partition Model* recently developed by Müller-Crepon, Schvitz and Cederman (2023), we obtain similar results.

Table 1: Covoting and shared mother tongue

| Dependent Variables: | Voting intention | | Party preference | |
|---|---|---|---|---|
| Model: | (1) | (2) | (3) | (4) |
| *Variables* | | | | |
| Shared mother tongue (0/1) | 0.178*** | 0.169*** | 0.172*** | 0.164*** |
| | (0.030) | (0.029) | (0.033) | (0.033) |
| Shared religion (0/1) | | 0.020* | | 0.020** |
| | | (0.008) | | (0.008) |
| Age similarity (decades) | | -0.004** | | -0.003 |
| | | (0.001) | | (0.002) |
| Shared gender (0/1) | | 0.000 | | -0.001 |
| | | (0.001) | | (0.001) |
| Shared education (0/1) | | 0.021*** | | 0.017*** |
| | | (0.004) | | (0.004) |
| Wealth similarity (sd) | | 0.004+ | | 0.003 |
| | | (0.002) | | (0.003) |
| Shared occupation (0/1) | | 0.028*** | | 0.025*** |
| | | (0.006) | | (0.006) |
| Geographic proximity (1'000km) | | 0.039* | | 0.039* |
| | | (0.019) | | (0.019) |
| Shared urban vs. rural (0/1) | | 0.025*** | | 0.023*** |
| | | (0.005) | | (0.006) |
| *Fixed-effects* | | | | |
| Country x Round | Yes | Yes | Yes | Yes |
| *Fit statistics* | | | | |
| Outcome mean | 0.460 | 0.460 | 0.458 | 0.458 |
| Countries | 26 | 26 | 28 | 28 |
| Respondents | 17,832 | 17,832 | 13,888 | 13,888 |
| Dyads | 6,506,927 | 6,506,927 | 3,815,492 | 3,815,492 |
| $R^2$ | 0.077 | 0.080 | 0.076 | 0.078 |
| Within $R^2$ | 0.021 | 0.024 | 0.020 | 0.022 |

*Standard errors clustered at the level of each respondent's mother tongue in parentheses.*
*Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1*

to 21.6 percent of coethnics among pairs of non-covoters.

At the macro-level, our results indicate that, on the margin, increasing ethnic homogeneity translates into a higher concentration of parties with an elasticity of .17 (see also Equation 7).[34] This positive elasticity stands in drastic contrast to the *negative* bivariate relationship observed when using country-level data (see Appendix A8), highlighting the caveats of ecological inferences drawn from aggregate data.[35]

We observe little systematic change in the aggregate effect of coethnicity on covoting intentions over time. When repeating our analysis for Afrobarometer rounds 3 to 7 in Figure 3 we find a slight upwards trend in the full sample, which includes increasingly many countries.[36] Yet, there is no significant increase in the effect of coethnicity once we subset the sample to countries that have always been surveyed. In other words, the upwards slope observed in the upper panels in Figure 3 results mostly from the increasing sampling of countries with more extensive ethnic voting.

In contrast, the estimated effects of coethnicity on covoting intentions vary within countries over time. We discuss three cases that feature prominently in previous studies, sometimes as examples that demonstrate the weakness or even absence of ethnic voting. Figure 4 displays results across Afrobarometer rounds 3-7 for Kenya, Malawi, and Mali. Political scientists typically describe Kenyan elections as classic cases of ethnic voting (Bratton and Kimenyi 2008). Our analysis supports this interpretation. Over time, the estimated effect of coethnicity increased on average, with a small decline in 2016 when shared ethnicity is associated with a 30 percentage point increase in covoting intentions, more than 1.5 times our main estimate.

The Kenyan illustration serves well to illuminate both the comparative advantages and remaining limitations of the CVR. The coethnicity coefficient $\beta_1$ can be

[34]The elasticity of the effective number of parties to changes in ethnic homogeneity depends on the value of other covariates, in particular the country fixed effects and the country-level of ethnic homogeneity.

[35]Our estimated country-level coethnicity effects correlate positively with Posner's (2004*a*) fractionalization index (Appendix Figure A10).

[36]For this inter-temporal analysis, we construct the coethnicity indicator based on respondents' self-identified ethnicity, which has been consistently asked since round 3.

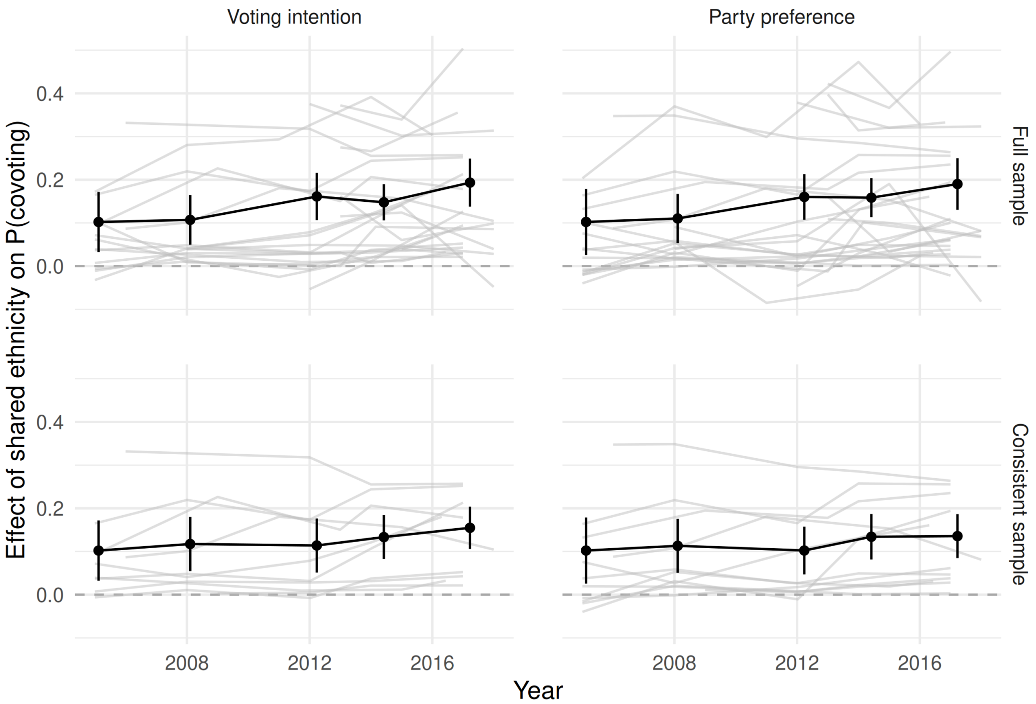



Figure 3: Effect over time, by Afrobarometer survey round

Note: Coefficients result from the fully specified model in Equation 10 estimated separately for each Afrobarometer survey round. The estimated models are thus equivalent to Models 2 and 4 in Table 1 except for the underlying sample and using respondents' self-identified ethnicity to construct the coethnicity indicator. 'Full sample' refers to all countries included in any one survey round, while 'consistent sample' refers to countries included since Afrobarometer Round 3. Grey lines plot country-by-country estimates over time, see Appendix Figures A21 to A23 for full results.

compared across survey rounds to reveal aggregate changes in the patterns of coethnig covoting conditional on covariates. It does not, however, reveal whether these aggregate changes result from individual changes in voting behavior, candidate entry and exit, or both. The right panel of Figure 4 shows the derived group- and party-based components of coethnic covoting. While their values can be affected by changes in other quantities besides coethnicity ($\beta_1$), including covoting ($\beta_0$) and covariate effects ($\gamma$), both measures run largely in parallel to $\beta_1$.[37] The results suggest that the Kenyan increase in coethnic covoting came with a greater increase in the homogeneity of candidates' voters as compared to a smaller increase in the homogeneity of groups' voting intentions.[38]

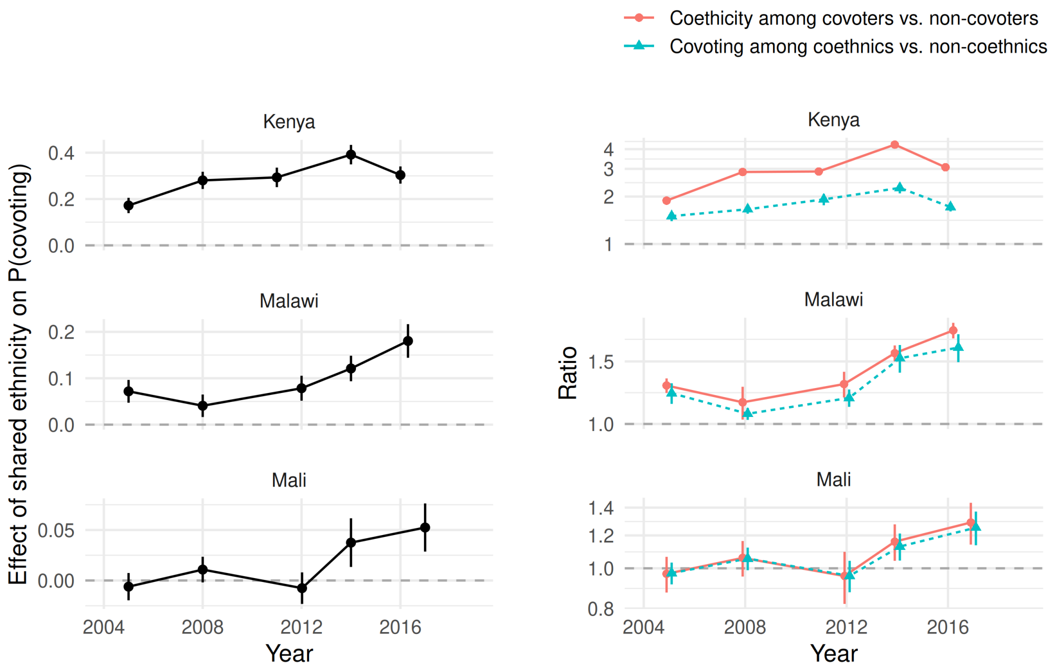
Next, we turn to Malawi, a country for which prominent studies diagnosed weak ethnic voting patterns. In line with Ferree and Horowitz' analysis of Malawi, we indeed observe relatively small coefficients for shared ethnicity in the run-up to the 2009 election, in which the ethno-regional voting "pattern broke down in dramatic fashion" (2010, 535). However, since then our results indicate a strengthening of the effect of coethnicity that reaches the estimated average effect for all elections in SSA in Afrobarometer round 7. Increasing voting differences between ethnic groups and rising ethnic differences between parties' voters run parallel to this trend. This development challenges recent work that identifies Malawi's persistent regional voting blocs and underlying shared economic interests as the better fitting explanation of covoting intentions (Boone et al. 2022).[39]

Finally, in a widely-cited study, Dunning and Harrison (2010, 21) "help explain why ethnicity has a relatively minor role in Mali ... [a] country in which ethnic identity is a poor predictor of vote choice." Our analysis confirms Dunning and Harrisons's verdict when they wrote their study in the late 2000s before the renewed outbreak of civil war in 2012 which also produced increasing communal

[37]Indeed, a simple regression across all country-rounds shows that the group- and party-based ratios together explain 92% of the overall variation in $\beta_1$ and 80% of its temporal variation within countries.

[38]Such changes can be driven by changing ethnic electoral coalition dynamics.

[39]Similarly, we estimate some of the strongest coethnicity effects on covoting intentions for Kenya (see above) and Zambia (Figure A23), Boone et al.'s other cases that serve to illustrate the influence of geography on voting blocs. In contrast to Boone et al. who impose a primacy of geographic blocs on the data, the CVR shows a relatively higher effect for shared ethnicity compared to geographic proximity in these cases.




(a) Effect of coethnicity on covoting

(b) Party- and group-based derivatives

Figure 4: Ethnic voting over time in Kenya, Malawi and Mali

Note: (a) Coefficients result from the fully specified model in Eq. 10 estimated separately for each Afrobarometer survey round since Round 3 with respondents' self-identified ethnicity as ethnicity indicator. The estimated models are thus equivalent to Models 2 in Table 1 in all respects except for the underlying sample. See Appendix Figures A21 to A23 for all countries in the sample.
(b) Ratios are computed by dividing the covariate-adjusted rate of coethnicity among covoters by the rate of coethnicity of non-covoters (red) and by dividing the rate of covoting among coethnics by the rate of covoting among non-coethnics (blue).

violence. Since then, the effect of shared ethnicity on Malian citizens' covoting intentions increased to a comparatively small effect of 6 percentage points in 2017. This development comes with more covoting within ethnic groups and heightened ethnic differences between parties' voter pools. We estimate similarly small effects of shared ethnicity in Burkina Faso and Senegal (see Figures A21 and A23), two other countries that Africanists have singled out as examples of weak ethnic voting (e.g., Koter 2013*a*). As in Mali, however, both cases see coethnics increasingly covote.

Together these trends underline the importance of broad comparative work like ours. It is important to note that changes in the effects of coethnicity on covoting over time and differences between countries can be driven by many factors not discernible from CVR estimates such as general changes in political preferences, party systems, or electoral institutions. Explicit analyses of effect heterogeneity can, however, shed light on such drivers as we show in exploratory analyses below.

### Robustness checks

We systematically test the robustness of our results to the omission of cross-cleavage interactions, the measure of ethnicity used, accounting for the potentially biasing effect of geography, changes in the setup of the data and estimation, as well as to using an alternative statistical network estimator. Our discussion below summarizes the results presented in Appendix F.

**Saturated interaction and multinomial regression models:** Our main explanatory variable "shared mother tongue" is based on an interaction between linguistic characteristics of respondents' $i$ and $j$ The main analysis might therefore be biased by the omission of substantively important cross-cleavage interactions, interactions between all characteristics (ethnicity, wealth, education, etc.) of voter $i$ with that of voter $j$ (Appendix A.4). As a remedy, we estimate fully interacted models.[40] We also compute the effects of coethnicity on covoting from multinomial choice models for each country in our sample. Appendix F.1 and F.2, respectively, show that the

[40] We obtain equivalent results when reducing the number of interactions via LASSO (Beiser-McGrath and Beiser-McGrath 2020).

results closely align with the estimates from our specification with simple control variables (Eq. 6).

**Accounting for potentially endogenous ethnicity:** Endogenous ethnicity (Pengl, Roessler and Rueda 2022) or identity misreporting (Green 2021) might bias our main estimates. In Appendix F.3, we therefore assess whether our results hold for respondents' current home language and ethnic self-identification, both of which are more malleable to change but potentially more precise. Supporting our main results, we find that a shared home language and ethnic self-identification increase covoting by 12 and 19 percentage points, respectively. We additionally estimate the effect of the pairwise linguistic proximity between respondents, finding that covoting increases with linguistic proximity. Varying the coarseness of language classification on the basis of the phylogenetic language tree yields a positive effect of coethnicity even if it is solely based on speaking a language belonging to the same of 6 language families. In combination, these results suggest that strategic misreporting or endogenous ethnic change are unlikely to substantively affect our results.

**Accounting for geographic variation:** Because many ethnic groups' feature regionally distinct "homelands", coethnic voting might simply emerge from an alignment of economic interests and political preferences of individuals from the same administrative region or even location (Boone et al. 2022; Boone 2024), geographies which might have partially shaped ethnic geography itself (Posner 2005; Müller-Crepon 2023).[41] Appendix F.4 shows that our results are robust to adding more fine-grained controls for geographic distance and joint residence in enumeration areas, districts and regions. These do not substantively interact with coethnicity. In addition, results from estimating our CVR only for the small, non-representative subset of respondent pairs residing in the same region or enumeration area show that coethnicity increases covoting in these selected pairs by 10 to 14 percentage points. This estimate decreases by another 2 to 4 percentage points once we include region and enumeration fixed effects, which soak up significant variation in

[41] Individuals' place of residence is, however, also endogenous to (ethnic) migration patterns.

ethnic and political homogeneity across locations. Overall, these results suggest that ethnic cleavages – while certainly influenced by geographic factors – produce strong electoral effects that are largely independent of geography.[42]

**Data construction:** We vary a number of choices made in the construction of our dyadic comparisons between survey respondents (see Appendix F.6). We first sequentially reduce the number of comparisons to the point of leaving only one comparison per respondent. This yields stable coefficient and uncertainty estimates. Second, we account for variation in the number of dyadic comparisons per country by weighting each dyad by the inverse number of dyads from its country such that every country receives the same weight.[43] This increases coefficient estimates slightly. Lastly, we recode preferences for "other" candidates and parties such that each such response is coded as its own candidate or party instead of being dropped. Doing so does not materially change the results.

**Model specification:** We furthermore test the robustness of our results regarding the most important modeling decisions (see Appendix F.7). We first reestimate the main specifications in Table 1 using logistic regression models, which yields equivalent results (Table A21).[44] Second, we test various ways of clustering our standard errors to account for the interdependence between dyadic comparisons, which yields less conservative estimates with the exception of clustering at the country level which yields marginally more conservative uncertainty estimates. Third, we implement different fixed effect specifications to account for potential sources of bias at the level of language groups, enumeration areas, and individual respondents on each side of a comparison. Doing so drastically improves model fit but does not substantively change the estimated effect of a shared mother tongue on covoting.

[42]We do note, however, that this finding is consistent with plausibly strong effects of economic geography on voting patterns (Boone 2024).

[43]This imbalance results from differing rates of missingness in the data.

[44]Similarly, estimating a network-based Probabilistic Partition Model (Müller-Crepon, Schvitz and Cederman 2023) yields consistent results (Appendix F.8).

## Mechanisms and heterogeneous effects

We can use the CVR to further test for mechanism and heterogeneity underlying the effect of coethnicity on covoting. Building on previous theoretical and empirical approaches at the micro-, meso-, and macro-level, the following exploratory, correlational analyses show strongest support for meso-level arguments about the importance of political mobilization of ethnic groups by "ethnic" parties. Evidence for micro-level arguments about ethnic biases and shared interest and macro-level theories on the importance of political institutions is mixed. While we focus on explaining cross-sectional heterogeneity, we note that similar analyses can be used to study drivers of temporal changes in the effect of coethnicity on covoting.

### Micro-level: Ethno-centrism and shared interests

Arguments in the micro-level literature can be divided into psychological explanations of ethnic voting that highlight preferential bias for coethnics (Horowitz 1985), and instrumental approaches that point to expectations of ethnic favoritism (Bates 1974), ethnicity as an information shortcut in low-information environments (Ferree 2006), as well as shared preferences among coethnics (Lieberman and McClendon 2013). We find no substantive moderating effect of ingroup bias proxied as respondents' ethnic rather than national identification (Appendix E.1). Expectations of favoritism are positively but insignificantly related to coethnic voting. Greater political knowledge and news consumption do not decrease the effect of coethnicity on covoting.[45]

Our test of the importance of ethnically structured interests proceeds in two steps. First, we extend the CVR to test whether coethnicity increases shared perceptions of pressing problems in a country and of governments' performance across 19 issue areas ranging from the economy to youth policy.[46] As Models 1 and 2 in Table 2 show, coethnicity comes with small, statistically significant increases in shared perceptions, yet these effects amount to less than 5 percent of the respective

[45]We continue to conservatively cluster standard errors at ethnic group level.

[46]For the former, we assess how much respondents' choice of the three most pressing problems overlap (0-1). For the latter, we code shared perceptions in a pair as 1 minus the average absolute rating difference (0-1) across issue areas (questions 56a–s).

outcome means. Second, we show that shared perceptions increase the likelihood of covoting between respondents. However, these effects do not diminish the estimated effect of coethnicity in Model 3. These results suggest that individual-level explanations do not explain the bulk of coethnic covoting, at least with the proxies available.[47]

Table 2: Shared mother tongue, perceptions

| Dependent Variables: | Problems | Performance | Voting intention |
|---|---|---|---|
| Model: | (1) | (2) | (3) |
| *Variables* | | | |
| Shared mother tongue (0/1) | 0.011*** | 0.015** | 0.167*** |
| | (0.002) | (0.005) | (0.029) |
| Problems | | | 0.013* |
| | | | (0.006) |
| Performance | | | 0.189*** |
| | | | (0.037) |
| *Fixed-effects* | | | |
| Country x Round | Yes | Yes | Yes |
| *Fit statistics* | | | |
| Controls | Yes | Yes | Yes |
| Outcome mean | 0.256 | 0.669 | 0.458 |
| Dyads | 6,397,649 | 6,466,584 | 6,361,704 |
| $R^2$ | 0.036 | 0.043 | 0.082 |
| Within $R^2$ | 0.006 | 0.010 | 0.027 |

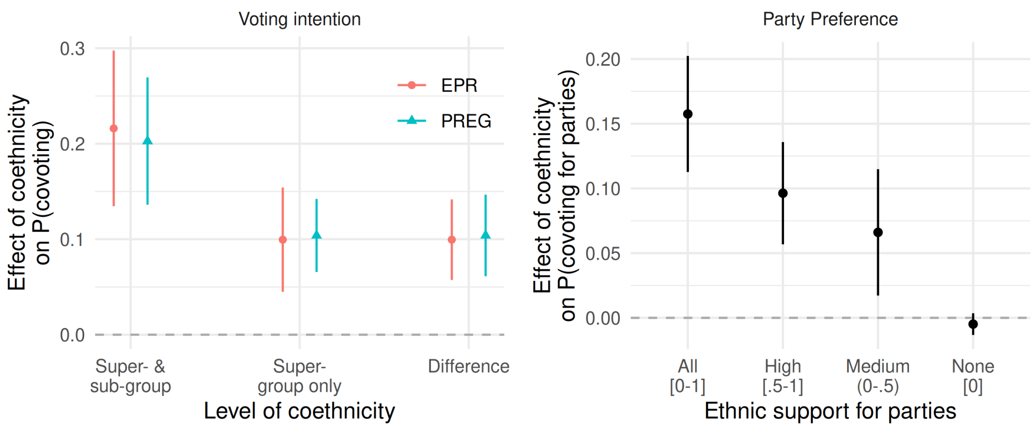
*Standard errors clustered at the level of each respondent's mother tongue in parentheses.*
*Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1*

### Meso-level: Politicized ethnic groups and ethnic parties

At the meso-level, we examine heterogeneity in the effect of coethnicity on covoting in two regards. First, we ask whether the political mobilization of (clusters of) ethnic groups as coded by Vogt et al. (2015) and Posner (2004*a*) increases covoting above and beyond the mere effect of speaking the same mother tongue. Suggesting substantive effects of ethnic groups' political relevance, Figure 5a shows that joint membership in a politically relevant group increases covoting by 10 percentage

[47] Unfortunately, the Afrobarometer does not include detailed batteries on policy ideal points which could improve the measurement of shared interests.

points, even if two respondents do not share the same mother tongue. Sharing the same mother tongue increases covoting by an additional 10 percentage points.




(a) Covoting within politically relevant "super-groups" from EPR and PREG

Note: Politically relevant "super-groups" from EPR (Vogt et al. 2015) and PREG (Posner 2004*a*) linked to mother tongues via Müller-Crepon, Pengl and Bormann (2022). Estimates from Tables A6 and A7.

(b) Covoting for parties with differing reliance on ethnic support groups

Note: Data on ethnic support groups from the VDEM V-Party data (Lührmann et al. 2020) linked to Afrobarometer Round 6 via Döring and Regel (2019). Estimates from Table A8.

Figure 5: Variation in the effect of coethnicity on covoting by groups' political relevance and parties' reliance on ethnic constituencies

Second, we disaggregate the coding of covoting to reflect whether respondents share a preference for a party that the V-Party data (Lührmann et al. 2020) codes as having a high or medium level of ethnic support, or none at all.[48] The results in Figure 5b suggest that two thirds of the overall effect of coethnicity on covoting is driven by covoting for parties with a high level of ethnic support, with the remainder being due to parties with a medium level of ethnic support. We find no substantive effect of coethnicity on covoting for parties with no ethnic support base. Their voter bases are spread uniformly across ethnic groups, which shows that our results align closely with experts' assessments of parties' core constituencies.

## Macro-level: Political institutions

Prior research highlights theoretical reasons to expect substantive variation between countries in the extent to which ethnic cleavages structure electoral politics

[48]The classification of parties derives from the share of V-Dem experts that agree that a party has an ethnic support group. We drop parties that are not included in the VDEM-Party data.

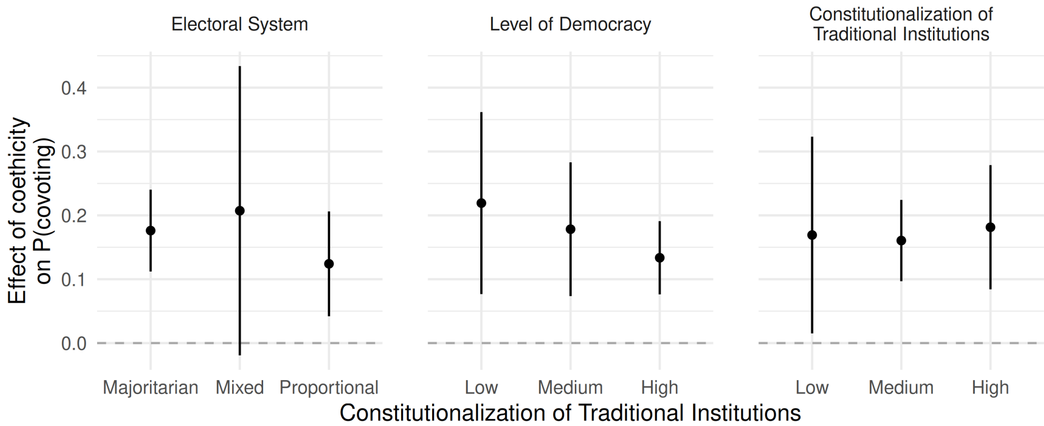



Figure 6: Effect of coethnicity on shared voting intentions by countries' institutional characteristics

Note: Coefficients result from the fully specified model in Eq. 10 estimated separately for countries with different electoral rules, levels of the V-Dem's Polyarchy index as well as constitutionalization of traditional institutions. See Appendix Tables A9 to A11 for underlying models.

(e.g. Huber 2012; Mozaffar, Scarritt and Galaich 2003). We analyze such heterogeneity along electoral systems, countries' level of democracy, and the strength of traditional institutions. We do not find substantively or statistically significant variation in the effect of coethnicity on covoting intentions across these three arguably relevant institutional dimensions (Figure 6 and Appendix E.3). While these findings shed doubt on the importance of institutional variation, our findings can only be understood descriptively as we do not account for potential endogeneity of the moderating factors. It is also possible that existing analyses find empirical support for theoretical mechanisms that operate at the meso and macro-levels. It is particularly important to note that institutions can shape strategic candidate entry. This would introduce differential selection bias into our analysis and bias estimated differences in the effect of coethnicity on covoting.

## Comparing cleavages

Finally, we compare the effect of coethnicity and that of other socio-economic similarities between respondents with respect to their rate of covoting intentions. To facilitate a fair comparison that takes account of differences between conditional and unconditional effects, Figure 7 plots the results of baseline models of the effect of each variable without any additional controls, as well as coefficient estimates from the fully specified models (see Table 1, Models 2 and 4).


Figure 7: Results by cleavage indicator

Note: Coefficient estimates from (1) baseline model that only include the respective variable and country-fixed effects equivalent to the setup of Models 1 and 3 in Table 1, and (2) fully specified models with controls (Eq. 10), equivalent to the setup in Models 2 and 4 in Table 1. Error bars denote 95% CIs.

Among identity cleavages, shared mother-tongues seem to be by far the strongest and most stable predictor of covoting. The first column in Figure 7 depicts our main results from Table 1. Next, effects associated with shared religion are positive but decrease once we condition on covariates. We presume that the unconditional effect of shared religion captures some of the effect of (correlated) sharing of mother-tongues. Across stated support for presidential candidates and legislative parties, we find no substantive effects of age and gender similarities. For age, we find a small *negative* effect of being close in age on covoting intentions, which suggests that party preferences within age groups are marginally more diverse than across them.

Economic similarities show substantively smaller effects on convergent voting intentions than those associated with shared mother tongues. Shared levels of education and occupation between respondents translate into an increase in the chance of supporting the same party by about 2 and 3.4 percentage points, respectively.

These effects are robustly estimated. Interestingly, proximity in wealth-levels between respondents does not significantly relate to covoting intentions between them.[49] This finding speaks to previous findings on economic voting in Sub-Sahara Africa (Bratton and Kimenyi 2008; Bratton, Bhavnani and Chen 2012).

Lastly, we find the geographic proximity correlates with shared support for presidential candidates and parties. In the unconditional baseline models, increasing geographic proximity by 1'000km comes with an increase in covoting intentions by 10 percentage points, consistent with the existence of regional voting blocs (Boone et al. 2022).[50] Yet, once we condition on all other cleavage measures, the effect of geographic proximity drops by half. This decrease supports the interpretation that geography correlates with voting preferences because of its reflection of economic incentives and ethnic identities. Shared urban or rural status has a consistent and statistically significant effect on covoting of approximately 2 percentage points when including controls. This is consistent with work on rural-urban cleavages on the continent (e.g. Harding 2010; Koter 2013*b*).

## Conclusion

In this paper, we introduce the Covoting Regression (CVR) model as a novel analytical approach to study the electoral effects of social cleavages in general and ethnic voting in Africa in particular. Shifting from individual support for coethnic candidates towards shared voting intentions between two individuals allows us to address two key methodological issues in existing work. First, we reduce selection bias that plagues micro-level studies when the supply of candidates does not allow survey respondents to express support for coethnic candidates, or forces them to do so in the absence of non-ethnic rival candidates (e.g., Ferree 2022). The CVR does not overcome selection bias that results from strategic choices of candidates to run, even if its consideration of ethnic bloc voting should reduce such bias in many contexts in comparison to previous approaches. Second, we avoid ecological inference inherent in meso and macro-level research that examines ethnic covot-

[49]The estimate is robust to expanding the set of variables used to construct the wealth index (Appendix F.5).

[50]Though note that our sample includes many small countries where such a change is unrealistic.

ing but fixes ethnic groups as the main unit of analysis while disregarding other cleavages. Coincidentally, we retain the advantages of micro and macro studies. The CVR model we introduce captures both individual-level effects, *and* recovers country-wide concentration indices such as the effective number of parties and the Herfindahl-Hirschman index of ethnic concentration. Finally, CVR operates at scale and enables broad comparisons over time and across countries without preventing country-specific insights.

Our empirical analysis of 28 countries and five survey rounds from the Afrobarometer indicates that language-based ethnicity continues to be the dominant electoral cleavage across Sub-Saharan Africa. The effect of coethnicity on voting intentions is at least five times larger than alternative cleavages including religion, shared urban or rural residence, geographic regions, as well as educational and occupational background. Although we find that coethnicity does not influence covoting equally across all survey rounds and countries, prominent case studies that question the effect of ethnicity in vote choices describe exceptions rather than broader trends across the African continent (Dunning and Harrison 2010; Ferree and Horowitz 2010; Boone et al. 2022). Finally, our analysis reveals the importance of group- and party-level characteristics in explaining the incidence of coethnic covoting, but yield no robust evidence for micro-level explanations around ethnocentric bias or policy perceptions, as well as macro-level arguments on the impact of electoral rules, democratic institutions, or traditional authorities (Huber 2012; Rabushka and Shepsle 1972; Baldwin 2013).

Our study opens up new avenues for studying ethnic and more generally cleavage-based political behavior in Sub-Saharan Africa and beyond. More precisely measured data on economic income, production patterns, and partisanship would enable us to gain deeper insight into class, economic, and psychological explanations of voting – three core concerns of voting research outside Africa. For example, incorporating local economic production data could be used to approximate residents' economic interests and model the extent to which they structure the formation of regional voting blocs that can span across ethnic divides (Boone 2024). While economic-instrumentalist and psychological factors have already received much attention in the study of ethnic voting, one major theory of voting, its sociological basis (Lazarsfeld, Berelson and Gaudet 1968/1944), has been widely overlooked by students of Sub-Saharan Africa (though see work on traditional institutions and norms, Baldwin 2013; Holzinger et al. 2019). Given appropriate data, the CVR model could, for example, test the effect of different social networks on voting by capturing the overlap in (the homogeneity of) social contacts.[51]

Beyond Sub-Saharan Africa, our analytical focus on covoting intentions lends itself to the study of the relative strength of different cleavages, such as the re-emergence of urban-rural divides (Cramer 2016), and the increasingly dominant nationalist-cosmopolitan division across western democracies (Kriesi et al. 2012). Our method might also benefit existing meso-level analyses of vote shares in small-scale spatial units, such as municipalities (Cagé and Piketty 2023). These analyses face similar challenges as the ones we discussed in the context of research on voting intentions in Sub-Saharan Africa. Rather than estimating the likelihood of covoting intentions at the individual-level, we would require compositional similarity scores between spatial units in terms of voting results as a function of similarities in their social structure. Lastly, our work shows that the CVR might benefit studies of the effects of cleavages on political attitudes and behavior more generally as it can accommodate more outcomes than just covoting, as for example shared (sets of) political attitudes and policy preferences between individuals. After all, social and political cleavages are an inherently *relational* concept and their effects on political preferences should be operationalized and studied as such.

[51]Spillover experiments constitute an attractive but more costly alternative methodology (e.g., Foos and De Rooij 2017).

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# Supplementary Material

## Table of Contents

# A Approaches to modeling ethnic voting

The following section contains formalized descriptions of modeling approaches to ethnic voting in existing work which we compare to the CVR. In the next section, we demonstrate how different approaches suffer from different biases with the help of Monte Carlo simulations.

## A.1 Ethnic Voting as Voter-Candidate/Party Identity Correspondence

Assume a setting with two parties $P$ ($A$, $B$) and two ethnic groups $G$ ($a$, $b$). Party $A$ runs candidates from group $a$ and party $B$ fields candidates from group $b$ (descriptive representation). Alternatively, parties $A$ and $B$ represent groups $a$ and $b$ respectively without fielding candidates of this ethnicity (substantive representation).

More generally, the predominant empirical approach assumes that any voter $i$ receives positive utility from voting for a party $p_i \in P$ that represents their group identity. Based on data on the set of parties $P_G \subset P$ that represent a given group $g \in G$, the probability of ethnic voting $EV$ in this approach is:

$$Pr(EV) = 1/N \sum_i Pr(p_i \in P_{g_i}|g_i) = 1 - 1/N \sum_i Pr(p_i \notin P_{g_i}|g_i) \quad \text{(A1)}$$

Importantly, to assess Eq. A1, it is necessary for researchers to specify the mapping between groups $G$ and parties (or candidates) $P$. The choices in this regard are shaped by important conceptual differences in approaches to ethnic voting. Huber (2012) distinguishes between a *group-based* perspective on ethnic voting and a *party-based* view. The former focuses on the degree to which each group votes homogeneously, suggesting few or only one party in each $P_G$ while not paying attention to any overlap between subsets. The latter, in turn, recognizes high levels of ethnic voting if all voters of a given party or candidate come from one group, thus letting party sets $P_G$ to be disjoint subsets of $P$.

In practice, the predominant approach of using voter-candidate coethnicity to operationalize ethnic voting falls into the party-based approach as each candidate is clearly attributed to only one group.[1] Empirical analyses of ethnic voting that rely on the voter-party correspondence approach thus reach the maximum estimate if all voters from one party come from one ethnic group, i.e., they rely on the party-based approach to ethnic voting. Specifically, they estimate lower levels of ethnic voting if a party draws support from multiple ethnic groups, even if all members of each group support the party. This is directly related to the selection biases highlighted by Ferree (2022) in cases where for some groups $g$ $P_g$ is an empty set or encompasses the full set $P$.

## A.2 Ethnic Voting as Coethnic Covoting

The covoting approach abstracts away from candidate/party identities or claims regarding group representation, and recasts ethnic voting as the tendency of voters

[1] Taking into account candidates' multiple identities could mitigate this slightly.

from the same group to vote for the same party. The probability of ethnic covoting ($ECV$) can be written as:

$$\begin{aligned}
Pr(ECV) =& \frac{1}{N(N-1)} \sum_{ij, i \neq j} Pr(p_i = p_j | g_i = g_j) + \\
& Pr(p_i \neq p_j | g_i \neq g_j) \\
=& 1 - \sum_{ij} Pr(p_i \neq p_j | g_i = g_j) - \\
& Pr(p_i = p_j | g_i \neq g_j)
\end{aligned} \tag{A2}$$

As a result, $Pr(ECV)$ is driven by both logics of ethnic voting. It increases as more coethnics vote for the same party (group-based logic) and as fewer non-coethnics find themselves voting together, which makes parties' more ethnically homogeneous (party-based logic).

More formally, we can use the main formulation of the CVR to derive two measures that directly reflect the group- and party-based logics of ethnic voting. First, we can compute the share of covoters among coethnics: If high, in particular high relative to the share of covoters among non-coethnics,[2] groups vote as blocks and Huber's group-based voting fractionalization (GVF) is high. The measure can be computed directly on the basis of the CVR as defined in Eq. 6:

$$\begin{aligned}
Pr(\text{covoting} \,|\, \text{coethnicity} = 1) &= \beta_0 + \beta_1 + \gamma \bar{\mathbf{x}} \\
Pr(\text{covoting} \,|\, \text{coethnicity} = 0) &= \beta_0 + \gamma \bar{\mathbf{x}}
\end{aligned} \tag{A3}$$

Following the party-based logic, we can compute the share of coethnics among covoters. If high, in particular if high relative the share of coethnics among those who do not covote,[3] parties feature ethnically homogeneous constituencies and Huber's party-based vote fractionalization (PVF) is high. Again, we can compute the measure directly from our model by scaling the previous concept by the relative rate of coethnicity compared to covoting:

$$\begin{aligned}
Pr(\text{coethnicity} \,|\, \text{covoting} = 1) &= \frac{\overline{\text{coethnicity}}}{\overline{\text{covoting}}} \, (\beta_0 + \beta_1 + \gamma \bar{\mathbf{x}}) \\
Pr(\text{coethnicity} \,|\, \text{covoting} = 0) &= \frac{\overline{\text{coethnicity}}}{1 - \overline{\text{covoting}}} \, (1 - (\beta_0 + \beta_1 + \gamma \bar{\mathbf{x}}))
\end{aligned} \tag{A4}$$

We illustrate the usefulness of the group- and party-based reformulation of the effect of coethnicity on covoting retrieved from a CVR.

**Scenario I:** A situation in which four groups (size .25 each) vote as blocks first for one party, and then for two parties (Figure A1). The CVR regression coefficient for coethnicity is 0 in the first situation as coethnics do not covote more than non-coethnics, but increases to .667 in the second situation where coethnics pairs covote

[2]This relative expression, e.g., through a ratio, is necessary to account for the base rate of covoting.

[3]This relative expression, e.g., through a ratio, is necessary to account for the base rate of coethnicity.

Situation 1

| | Groups | | | |
|---|---|---|---|---|
| | A | B | C | D |
| Party 1 | 1 | 1 | 1 | 1 |

$\Longrightarrow$

Situation 2

| | Groups | | | |
|---|---|---|---|---|
| | A | B | C | D |
| Party 1 | 1 | 1 | 0 | 0 |
| Party 2 | 0 | 0 | 1 | 1 |

Figure A1: Scenario I

Situation 1

| | Groups | |
|---|---|---|
| | A | B |
| Party 1 | 1 | 0 |
| Party 2 | 0 | 0.5 |
| Party 3 | 0 | 0.5 |

$\Longrightarrow$

Situation 2

| | Groups | |
|---|---|---|
| | A | B |
| Party 1 | 1 | 0 |
| Party 2 | 0 | 1 |

Figure A2: Scenario II

more often than non-coethnic ones: $Pr(covoting|coethnicity = 0)$ decreases from 1 to 33 percent. Across this scenario, coethnics always covote ($Pr(covoting|coethnicity = 1) = 1$ and $Pr(coethnicity|covoting = 0) = 0$). The coethnicity coefficient changes as a function of the increasing ethnic homogeneity of parties' voters: the chance that two covoters are coethnics $Pr(coethnicity|covoting = 1)$ increases from 25% to 50%.[4]

**Scenario II:** A scenario in which two groups (size .5 each) vote first for three purely ethnic parties whereby group B splits its vote, and then for two mono-ethnic parties (Figure A2). The CVR regression coefficient for coethnicity is 0.75 in the first situation, and increases to 1 in the second as coethnics always covote. Non-coethnics never covote in either situation ($Pr(covoting|coethnicity = 0) = 0$. Covoters are always coethnics ($Pr(coethnicity|covoting = 1) = 1$), so that the change in the coefficient of coethnicity is driven by the decreasing political fractionalization of group B: the chance that two coethnics covote $Pr(covoting|coethnicity = 1)$ increases from 75% to 100%.[5]

**Scenario III:** In a third scenario, we compare a first situation in which two groups of equal size each split their votes into two monoethnic parties with a second situation in which four groups vote as homogeneous blocks for two parties (see Figure A3). Here, the CVR estimate for the effect of coethnicity on covoting increases from .5 in situation 1 to .667 in situation 2. This change is the product of the increase of the chance that coethnics covote $Pr(covoting|coethnicity = 1)$ from 50% to 100% and the decrease in the chance that covoters are coethnics $Pr(coethnicity|covoting =$

[4]Both of Huber's indices, GVF and PVF, take values of 0 in situation 1 because ethnic groups do not differ in their choice of the one party. GVF increases to 0.5 in situation 2 while PVF increases to .35.

[5]Huber's GVF increases from .433 in situations 1 to .5 in situation 2. The PVF index remains at .5 for both situations as parties maintain their average level of ethnic distinctiveness.

Situation 1

| | Groups | |
|---|---|---|
| | A | B |
| Party 1 | .5 | 0 |
| Party 2 | .5 | 0 |
| Party 3 | 0 | .5 |
| Party 4 | 0 | .5 |

$\Longrightarrow$

Situation 2

| | Groups | | | |
|---|---|---|---|---|
| | A | B | C | D |
| Party 1 | 1 | 1 | 0 | 0 |
| Party 2 | 0 | 0 | 1 | 1 |

Figure A3: Scenario III

1) from 100% to 50%.[6]

## A.3 Voter-Candidate Correspondence versus covoting

If $i, j$ are iid then it is possible to express ethnic covoting expressions in voter-candidate correspondence terms. In our initial example, $Pr(AA|aa) = Pr(A_i|a_i) * Pr(A_j|a_j)$. Similarly, $Pr(AA|bb) = Pr(A_i|b_i) * Pr(A_j|b_j)$. While it is thus principally possible to link voter/candidate correspondence to the covoting approach, it also becomes clear that the two ways of conceptualizing ethnic voting differ. Whereas, for example, $Pr(AA|aa)$ and its constitutive terms $Pr(A_i|a_i)$ and $Pr(A_j|a_j)$ contribute positively to both $Pr(EV)$ and $Pr(ECV)$, $Pr(AA|bb)$ and its constitutive terms $Pr(A_i|b_i)$ and $Pr(A_j|b_j)$ contribute negatively to the probability of ethnic voting but positively to the probability of ethnic covoting (see Eqs. A1 and A2).

This divergence is directly related to the fact that the party-coding in the example prioritizes the party-based approach of ethnic voting over the group-based approach. In contrast, the CVR model captures both group and party-based conceptualizations of ethnic voting. It increases as supporters for one party come increasingly from the group that the party represents ($Pr(AA|aa)$ or $Pr(BB|bb)$). Similarly, the shared ethnicity coefficient in the CVR model increases if members from one group increasingly support one party, even if that party does not represent that group via coethnics ($Pr(AA|bb)$ or $Pr(BB|aa)$). In contrast, ethnic voting would decrease in the voter-party correspondence approach. This is the source of selection bias that arises if voters of a given party do not have coethnics/a party of their own ethnicity on the ballot.

## A.4 Individual choice models and the CVR

Most generally, it is possible to reformulate the CVR by bringing together individuals' candidate choice probabilities as typically modelled via a multinomial linear (or logistic) regression:

$$P(party_{p,i}|X_i) = \delta_{0,p} + \theta_p \mathbf{z}_i + \epsilon_{p,i} \tag{A5}$$

where individuals probabilities to vote for parties $party_p \in P$ is affected in a party-specific way by a constant $\delta_{0,p}$ and covariates $\mathbf{z}_i$ scaled by coefficient vector $\theta_p$. The

[6] Huber's GVF increases from .352 to .5 between situations 1 and 2 as groups become more homogeneous in their political preferences. In parallel, PVF decreases from .5 to .353 as parties' voter base become more ethnically diverse.

error $\epsilon_{p,i}$ is in expectation 0 and assumed to be iid. A categorical coding of $i$'s ethnic group among the covariates $\mathbf{z}_i$ can be used to estimate ethnic voting as the effect of a given ethnic identity on the probability of voting for a given party, thus combining a party- and group-centric understanding. A groups' voting homogeneity can be expressed in terms of its deviation from overall vote choices. A party's ethnic makeup is in turn reflected by the extent to which its voters come disproportionally often from a subset of ethnic groups.

Using Eq. A5, we can reformulate the main CVR (Eq. 6) as the sum of the joint probability that individuals $j$ and $i$ together vote for any party $p \in P$:

$$Pr(covote_{i,j}|\mathbf{z}_i, \mathbf{z}_j) = \sum_{p \in P} Pr(p_i|\mathbf{z}_i)\, Pr(p_j|\mathbf{z}_j) \tag{A6}$$

This formulation can be used to convert the results of a multinomial choice model into covoting probabilities and the effects of cleavages on them. Specifically, we can estimate Eq. A5, generate predicted probabilites at specific values of a predictor of interest such as ethnicity while keeping all other covariates at their sample mean (or any other arbitrary value). Using these predicted probabilities we can, in a second step, compute the overall covariate-adjusted effect of coethnicity on covoting ($\sim \beta_1$ in Eq. 6) by comparing how likely it is that two coethnics covote as compared to two non-coethnics. We use this approach as a robustness check in Appendix F.2.

Combining Eqs. A5 and A6 yields additional insights as to the potential importance of considering cross-cleavage interactions in addition to the main CVR's segregated treatment of cleavages. In particular, Eq. A6 can be reformulated as

$$= \sum_{p \in P} \big((\delta_{0,p} + \theta_p \mathbf{z}_i)(\delta_{0,p} + \theta_p \mathbf{z}_j)\big) \tag{A7}$$

$$= \sum_{p \in P} (\delta_{0,p}^2 + \delta_{0,p}\theta_p(\mathbf{z}_i + \mathbf{z}_j) + (\theta_p \cdot \theta_p^\intercal) \odot (\mathbf{z}_i \cdot \mathbf{z}_j^\intercal)) \tag{A8}$$

$$= \sum_{p \in P} \delta_{0,p}^2 + \sum_{p \in P} \delta_{0,p}\theta_p(\mathbf{z}_i + \mathbf{z}_j) + \sum_{p \in P} (\theta_p \cdot \theta_p^\intercal) \odot (\mathbf{z}_i \cdot \mathbf{z}_j^\intercal) \tag{A9}$$

Comparing Eq. A9 with the main CVR in Eq. 6 suggests several ways in which its "short" formulation might have to be extended to capture biases originating at the level of individuals' vote choice. In particular, (1) $\sum_{p \in P} \delta_{0,p}\theta_p(\mathbf{z}_i + \mathbf{z}_j)$ suggests adding the sum of covariates $\mathbf{z}_i$ and $\mathbf{z}_j$ to the model, essentially capturing overall covoting rates among certain groups of individuals. The respective coefficient estimates in the CVR would then capture the sum of party-specific coefficients $\sum_{p \in P} \delta_{0,p}\theta_p$.[7]

Adding more complexity, (2) the term $\sum_{p \in P} (\theta_p \cdot \theta_p^\intercal) \odot (\mathbf{z}_i \cdot \mathbf{z}_j^\intercal)$ denotes one additional term for each interaction between each covariate in $\mathbf{z}_i$ and each covariate in $\mathbf{z}_j$. In our case of ethnic cleavages, that includes not only all interactions between all ethnic groups of individuals $i$ and $j$, but also the interaction of $i$'s ethnicity with all other cleavage measures for $j$ (urban, wealth, etc.), as well as all interactions among these between $i$ and $j$, and vise-versa. Again, the respective coefficient

[7] The most straightforward and robust way to do so is to add fixed effects for individuals $i$ and $j$, see Appendix F.

estimates for the cross-cleavage interactions in the CVR would then capture the sum of party-specific coefficients $\sum_{p \in P}(\theta_p \cdot \theta_p^{\intercal})$.[8]

Doing so can be necessary because our measure of pair-wise coethnicity captures the interaction of the ethnicity of $i$ and $j$. Put differently, it is the mean effect across all interactions where $i$ and $j$'s ethnic identity is the same. To prevent omitted variable bias from affecting the estimate, it is thus potentially necessary to interact the respective control variables with the main cleavage of interest as well as interactions with each other (Beiser-McGrath and Beiser-McGrath 2020). Empirically, this requires the estimation of a large set of interaction terms, in particular where cleavages are based on categorical variables such as ethnicity, religion or occupation. In turn, problems from overfitting and loss of efficiency might result. Using models that identify relevant interactions from the data is one strategy to tackle these issues (see Beiser-McGrath and Beiser-McGrath 2020, for a discussion). We implement such models for a dedicated robustness check discussed in Appendix F.1.

# B Monte Carlo Simulations: Selection and ecological inference biases

Based on our discussions in Appendix A, this section presents two sets of Monte Carlo simulations that address (1) selection bias occurring in voter-candidate correspondence models compared to the CVR and (2) omitted variable bias that might affect the CVR and leads to biases from ecological inferences in meso–level approaches.

## B.1 Simulating Selection Bias

We simulate two types of selection biases that can affect the voter-correspondence and CVR model. The first consists in selection bias the occurs when researchers code some voting – typically for a coethnic candidate or party – as "ethnic" compared to non-ethnic votes cast for other parties. The second type of selection bias originates in endogenous party entry and exit, which might be causally related to the extent of ethnic voting in one or several ethnic groups. Our simulation shows that the correspondence model is affected by both biases, while the CVR does not incur biases from endogenous coding decisions but is affected by endogenous party exit and entry.

Our simulations presume 3 ethnic groups a, b, and c, each representing 1/3 of the population. Groups a and b exhibit fixed voting behavior towards two parties, A and B: 40% of their members vote "non-ethnically", splitting their vote equally between the two parties.[9] 60% of their members have an "ethnic" preference, voting for their "ethnic" party (a for A and b for B). As a result, members of a cast 80% (20%) of their votes for A (B), and members of b vote for B at a rate of 80%

[8] Because dyads are in principle undirected, we can force interactions to have the same coefficient irrespective of the order of $i$ and $j$ in the data.

[9] This could be due to a second substantive axis of party competition, a cross-cutting cleavage, or even simple indifference between the two parties.

(20%).[10] The third group, c, varies in the degree to which its members vote ethnically between 0 and 100%, thus varying the "true" value of aggregate ethnic voting alongside. Non-ethnic voters in c always split their votes equally between parties A and B.

In a first set of simulations (Scenario 1), ethnic voters from group c vote for a third party C. In the second set (Scenario 2), party C does not exist and ethnic voters from c vote for their "second-best" choice, party B. Here, we assume that B is not regarded as ethnically representative of group c by the researcher following the voter-correspondence approach. Non-ethnic voters from a, b, and c never vote for party C, a choice which is conservative but clarifies the contrast between the scenarios.[11]

For both scenarios, we generate 1000 draws of 1000 "survey respondents" and sample their party choices. We then estimate the average level of ethnic voting according to the voter correspondence model and the effect of coethnicity on covoting in the CVR. In Scenario 2, which does feature party C, we simulate two more researcher choices in the correspondence approach: for one, keeping group c in the analysis (with all members voting non-ethnically); for another, dropping group c due to the inability of voting ethnically.

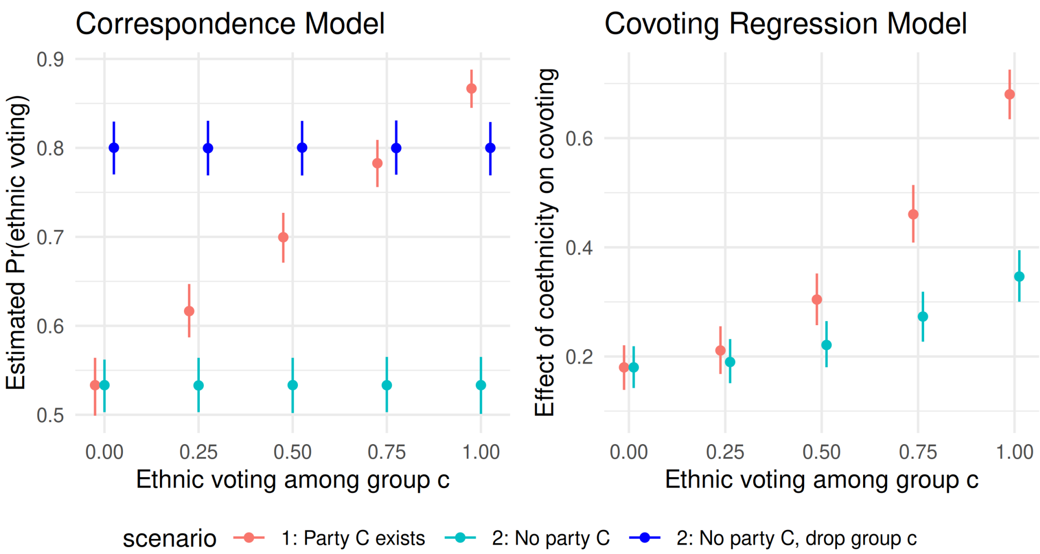



Figure A4: Selection bias in the voter-correspondence model and CVR

Figure A4 presents the results. Recall that the true extent of aggregate ethnic voting varies with the increasing share of members of group c to vote ethnically. In the left panel and for baseline Scenario 1 (red point ranges), we see that the correspondence approach clearly captures the increase of average ethnic voting by 33 percentage points. This is caused by the increasing share of ethnic voters among group c, which party C unambiguously separates from all other voters.

[10] Although only 60% of group members in a and b have an ethnic preference, we obtain an overall rate of 80% voting for their respective party. The discrepancy arises because half of the 40% of members without any ethnic preference split their vote equally between parties A and B.

[11] The entry of C would typically attract at least some "non-ethnic" voters from all groups which would muddy the comparisons.

In Scenario 2, that drops party C, the average rate of ethnic voting in the correspondence approach does not change, regardless of whether we keep (light blue) or drop group c (dark blue). Recall though that group c increases its vote for the "second-best" party B. Keeping group c in the analysis increasingly underestimates ethnic voting as the share of members of c would like to vote ethnically increases. The reason is that correspondence approaches do not consider group-based conceptualizations of ethnic voting. Dropping group c from the analysis overestimates ethnic voting for most parameter values, in particular when fewer members of group c prefer to vote ethnically.

Inportantly, we find that the correspondence approach overestimates the extent of ethnic voting by the share of non-ethnic voters that vote for "their" ethnic party due to chance. Even when dropping group c, we thus overestimate ethnic voting among groups a and b by 20 percentage points at 80 instead of 60 percent. This is because *all voters* for the ethnically corresponding party of a group are coded as "ethnic voters," even those 40% who randomly split their vote between parties A and B. Similarly, the base-rate of ethnic voting in Scenario 1 without ethnic voting by c is overestimated: it should be at 40 instead of 53.3 percent.

Turning to the CVR, the model clearly captures the increase in ethnic voting – expressed as the effect of coethnicity on covoting – in Scenario 1 (red point ranges). When no member of c votes ethnically, the effect of coethnicity on covoting is .18. This is exactly the result expected from the setup, as ethnic voters among a and b disproportionally vote for their parties A and B. The CVR correctly accounts for the covoting among coethnics that occurs due to pure randomness among the non-ethnic voters across all groups which split their votes across A and B. As a result, coethnics covote at a rate of 62 percent, and non-coethnics at a rate of 44 percent, leaving an effect of coethnicity of .18. As expected, this effect increases to a maximum of .68 as members of c shift their votes towards mono-ethnic party C, increasing covoting among coethnics from 62 to 78 percent, while strongly decreasing covoting among non-coethnics to 10 percent.

In Scenario 2, where party C does not exist, the effect of coethnicity on covoting increases with greater ethnic voting among c, but in a less pronounced manner. In this scenario, the rate of covoting among non-coethnics remains constant as voters c shift their votes from A to B. In contrast, however, the rate at which coethnics covote increases strongly from 62 to 78 percent, just as much as in Scenario 1, as the vote share for B increases in group c.

These results show that the CVR avoids the selection bias that comes with the correspondence approach when not all groups have the choice to vote for or against a coethnic. It also accounts for the baseline rate at which we would expect groups to vote for each party without any ethnic voting. Lastly, the coethnicity estimate in the CVR model is not robust to selection from potentially endogenous party entry and exit which remains an issue common to all observational approaches to studying ethnic voting.[12] However, the bias we observe is smaller than in the correspondence approach as the CVR captures the group-based logic of ethnic voting which remains unaffected by the entry/exit of party C. Moreover, assuming

[12]Note that the endogeneous party entry of the sort modeled here would also affect Huber (2012) measures of ethnic voting. Moving from Scenario 1 to 2 at positive levels of ethnic voting in group c would decrease party vote fractionalization as well as group vote fractionalization.

sincere voting and perfectly proportional electoral rules, the bias is largest when high demand of ethnic voting among members of c makes it least likely that no coethnic party would run. This claim rests on other factors not influencing voting decisions. The size of the bias likely varies with country-specific group sizes, the presence or absence of intra-group divisions, and individual candidate characteristics. In plurality or majoritarian electoral systems that incentivize strategic voting, the CVR would most likely suffer from selection bias even if members of group c have strong preferences for coethnic candidates if the latter do not run (or are not voted for) for strategic reasons.

### B.2 Simulating the impact of omitted variables and the threat of ecological inference

While group-based measures of ethnic voting, like those proposed by Huber (2012), reduce the problem of selection bias, they potentially create new challenges in the context of multiple cleavages that affect voting decisions. When applying group-based measures, researchers decide that one cleavage is the primary cleavage in their analysis. This may mask relevant political dynamics if this cleavage is correlated with another cleavage dimension, which also decisively influences electoral politics.

We simulate such a situation in a setting with two parties ($A$, $B$) and two ethnic groups ($1$, $2$), with a size of .4 and .6, respectively. However, voters have an additional urban/rural status with a mean of .5 and varying degrees of correlation (0 to .9) with their group identity. We model the probability of voting for party B in this Scenario 1 as:

$$Pr(VoteB_i) = .2 + .3\,\mathbb{1}(group_i = 2) + .2\,urban \tag{A10}$$

Table A1: Monte Carlo specifications for four different scenarios to test the influence of omitted confounders/cleavage dimensions.

| | Scenario | | | |
|---|---|---|---|---|
| | (1) | (2) | (3) | (4) |
| Parties | 2 | 2 | 2 | 4 |
| Ethnic Groups | 2 | 4 | 4 | 4 |
| $\beta_{Ethnicity}$ | 0.3 (Group 2) | 0.3 (2&4) | 0.3 (2&4) | 0.15 (Party B), 0.15 (C), -0.1 (D) |
| Urban/Rural | Yes (0.5/0.5) | Yes (0.5/0.5) | Yes (0.5/0.5) | Yes (0.5/0.5) |
| $\beta_{Urban}$ | 0.2 | 0.2 | 0.2 | 0.1 (Party B), 0.05 (C), 0 (D) |
| Income | No | No | $\mathcal{N}(0.5, 0.2)$ | $\mathcal{N}(0.5, 0.2)$ |
| $\beta_{Income}$ | | | 0.1 | 0.1 (Party B), 0.05 (C), 0 (D) |

We extend this setup in three steps to more realistic scenarios. In Scenario 2, we add two more groups, which simply mirror groups 1 and 2, such that group 3 behaves as $1$ and group $4$ as $2$ does in their voting behaviors and urban/rural status.[13] In Scenario 3, we add a second alternative cleavage, say income, which takes on continuous values (normally distributed, mean = .5, sd = .2), increases voting for party B with a coefficient of .1, and correlates with groups to the same extent as urban/rural status. The correlation between urban/rural status and income increases with the one with group identities. In Scenario 4, we add two more parties $C$ and $D$, and adjust coefficients in Eq. A10 downwards to avoid negative probabilities.[14] Table A1 summarizes the key parameters across the four scenarios

For each scenario, we vary the correlation between group identity and the alternative cleavages in 10 steps between 0 and .9. For each step and scenario, we simulate 1,000 voters 1,000 times. We evaluate the scenarios in three ways. We first compute Huber's (2016) Group Vote Fractionalization index, using ethnic groups and urban/rural status as group determinants. Second, we estimate the CVR models without covariates, with simple controls (joint urban/rural status, income difference), and a fully saturated model with all cross-cleavage interactions (see Appendix A.4). Finally, we show results that aggregate group-specific coefficients from party-specific individual-level analysis of the form of Eq. A10 with and without covariates. As explained in Appendix A.4, we use group- and party-specific predicted voting probabilities at the mean of all covariates to derive the overall, covariate-adjusted difference between covoting among coethnics and non-coethnics.

Figure A5 plots the values of Huber's group vote fractionalization (GVF) computed separately for ethnic groups and urban/rural strata. While we recuperate the true value of ethnic (and urban/rural) GVF (red dotted line) across all scenarios at zero correlation between ethnic groups and the alternative cleavages, both GVF measures increase as the correlation rises, even though the effect of each cleavage on vote choice remains constant by construction. The reason for the overestimation derives from the omitted impact of individuals' urban status (ethnicity) that increasingly feeds into the ethnicity (urban status) estimate with their growing correlation.

Turning to the results of the CVR and individual-level models in Figure A6 we find a similarly over-estimated effect of coethnicity on covoting in models without covariates at increasing levels of correlation between ethnicity and alternative cleavages (red line). Again, this upward bias stems from the impact of the omitted alternative cleavages (rural/urban status and income). In the CVR, this bias can be partially mitigated by adding "simple controls" for shared rural-urban status and income differences between individuals (turquoise line). Once we add all interactions between the characteristics of individuals' $i$ and $j$ in a pair, the CVR recuperates the true effect of coethnicity on covoting, just as the individual-level model does (blue line).

These results show that the CVR provides an unbiased estimate of the effect of coethnicity on covoting when confounders are observable and correctly specified. While the "simple" specification that models cleavages separately can be indicative of the existence of bias when comparing it to a no controls model, it does not necessarily capture all biases to their full extent. To that intent, researchers should

[13]We adjust group sizes of groups 1 to 4 to .2, .3, .1, .4.

[14]Party A is the baseline category. Parties B, C, D have a constant $\beta_0$ of .1, .1, .15; effects of groups 2 and 4 of .15, .15, -.1; an effect of urban status of .1, .05, 0; and of income of .1, .05, 0, respectively.

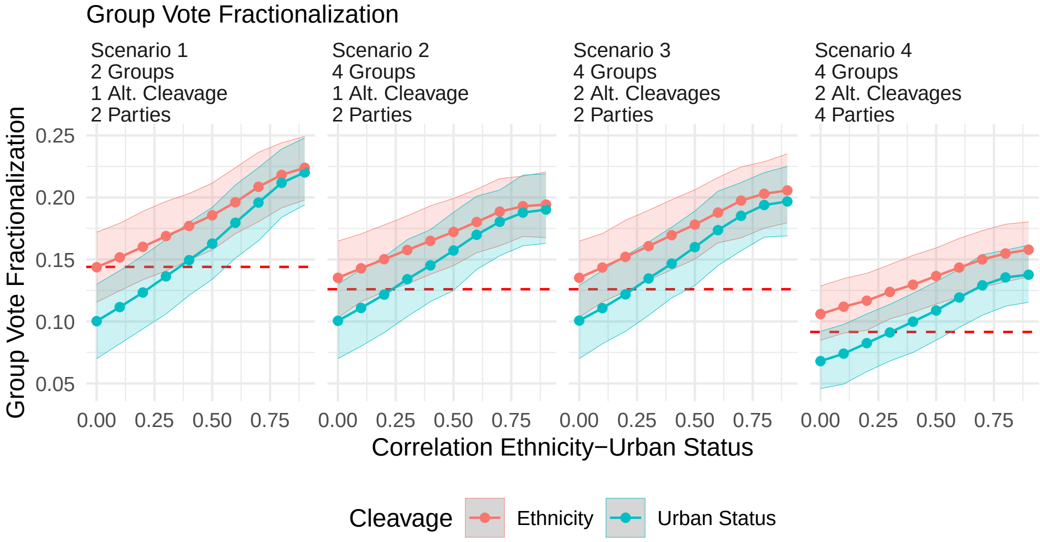



Figure A5: Monte Carlo simulation results: Huber's Group Vote Fractionalization. Ground truth value as red dotted line.

model a fully specified interactive model, which can be reduced to an empirically relevant set of interaction using, e.g., LASSO (see below in Appendix F.1). The results also show that conditional covoting probabilities can be successfully derived from individual-level models as explained in Appendix A.4 though the derivation of correct standard errors proves challenging.[15]

[15]This is because the results depend on separate models for each party.

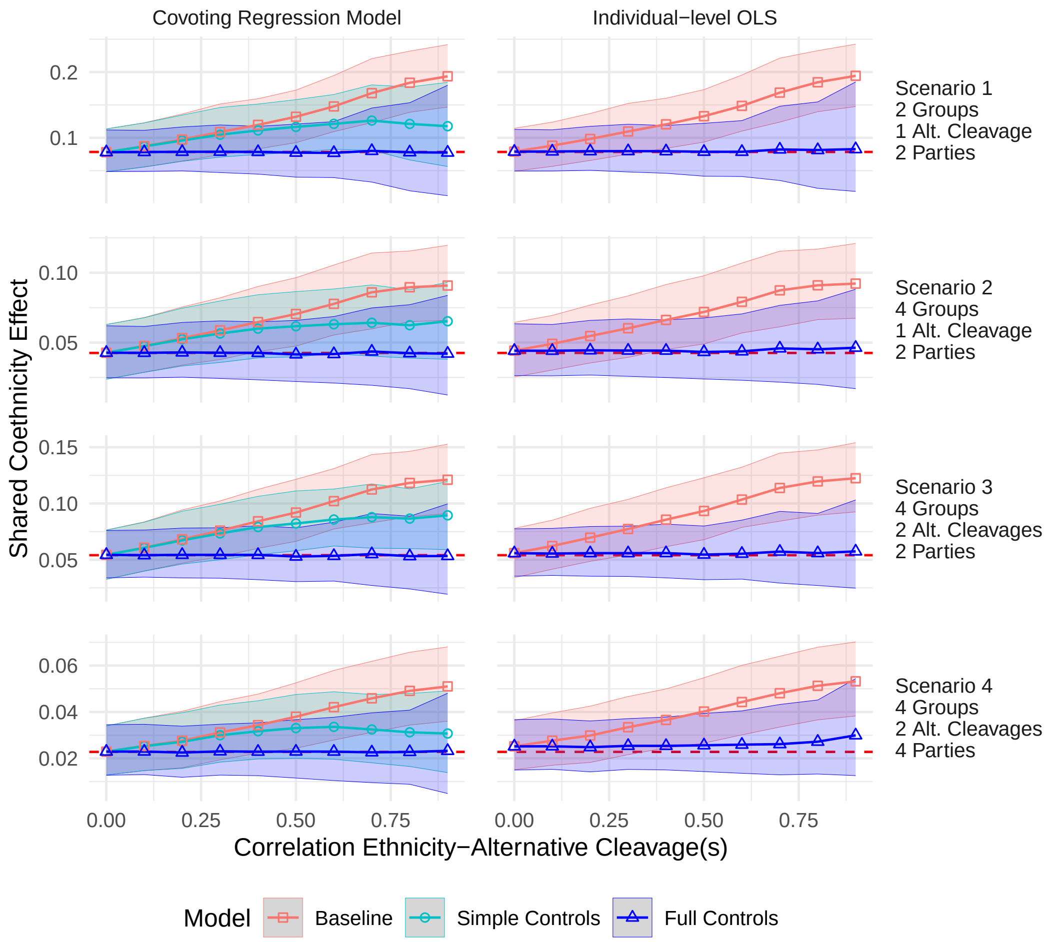



Figure A6: Monte Carlo simulation results: CVR and individual-level approach. Ground truth effects as red dotted line.

# C Summary statistics

Table A2: Summary statistics of variables in presidential voting intention data.

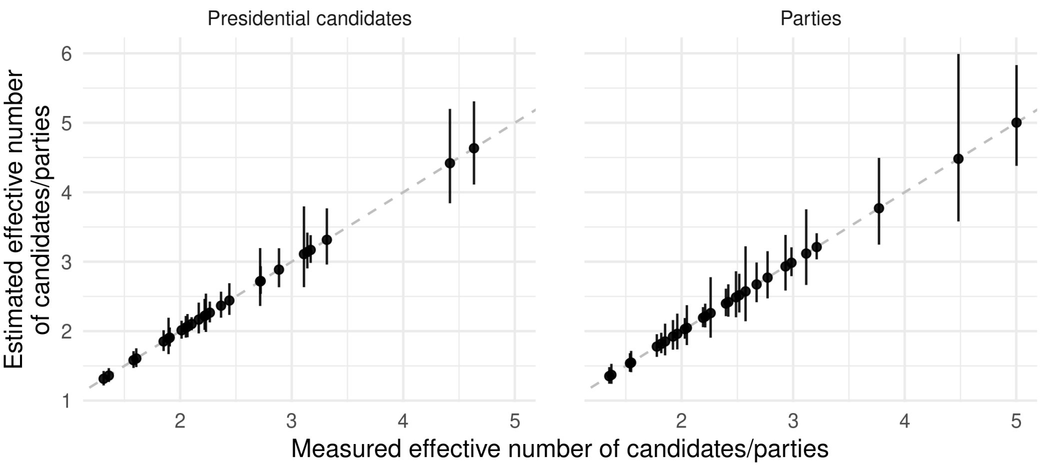
| Statistic | N | Mean | St. Dev. | Min | Max |
|---|---|---|---|---|---|
| Voting intention | 6,506,927 | 0.460 | 0.498 | 0 | 1 |
| Shared mother tongue | 6,506,927 | 0.216 | 0.412 | 0 | 1 |
| Shared religion | 6,506,927 | 0.334 | 0.472 | 0 | 1 |
| Age similarity | 6,506,927 | −1.573 | 1.282 | −8.800 | 0.000 |
| Shared gender | 6,506,927 | −0.500 | 0.500 | −1 | 0 |
| Shared education | 6,506,927 | 0.370 | 0.483 | 0 | 1 |
| Wealth similarity | 6,506,927 | −1.038 | 0.790 | −4.131 | 0.000 |
| Shared occupation | 6,506,927 | 0.201 | 0.401 | 0 | 1 |
| Geographic proximity | 6,506,927 | −0.328 | 0.286 | −2.359 | 0.000 |
| Shared urban/rural | 6,506,927 | −0.433 | 0.495 | −1 | 0 |

Table A3: Summary statistics of variables in party preference data.

| Statistic | N | Mean | St. Dev. | Min | Max |
|---|---|---|---|---|---|
| Party preference | 3,815,492 | 0.458 | 0.498 | 0 | 1 |
| Shared mother tongue | 3,815,492 | 0.276 | 0.447 | 0 | 1 |
| Shared religion | 3,815,492 | 0.340 | 0.474 | 0 | 1 |
| Age similarity | 3,815,492 | −1.643 | 1.317 | −8.800 | 0.000 |
| Shared gender | 3,815,492 | −0.496 | 0.500 | −1 | 0 |
| Shared education | 3,815,492 | 0.371 | 0.483 | 0 | 1 |
| Wealth similarity | 3,815,492 | −1.030 | 0.781 | −4.408 | 0.000 |
| Shared occupation | 3,815,492 | 0.209 | 0.407 | 0 | 1 |
| Geographic proximity | 3,815,492 | −0.309 | 0.268 | −2.359 | 0.000 |
| Shared urban/rural | 3,815,492 | −0.419 | 0.493 | −1 | 0 |

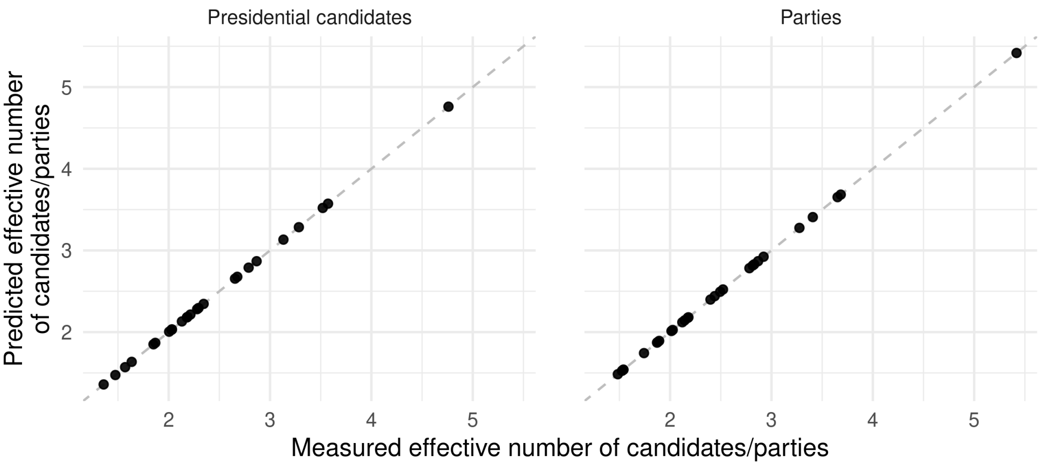
# D Empirical relationship to macro-, meso-, and micro-level approaches

## D.1 Macro-level: Effective number of parties and ethnic fractionalization




(a) Estimated and measured effective number of parties

Note: Estimates are derived as the inverse intercept of an otherwise empty regression model estimated separately for each country, standard errors are clustered on the level of individuals.




(b) Predicted and measured effective number of parties

Note: Predicted ENP is derived as the inverse average fitted probability of co-voting obtained from the fully specified regression model estimated separately for each country.

Figure A7: Empirical relation between measured Effective Number of Parties and the Linear Probability Model of co-voting

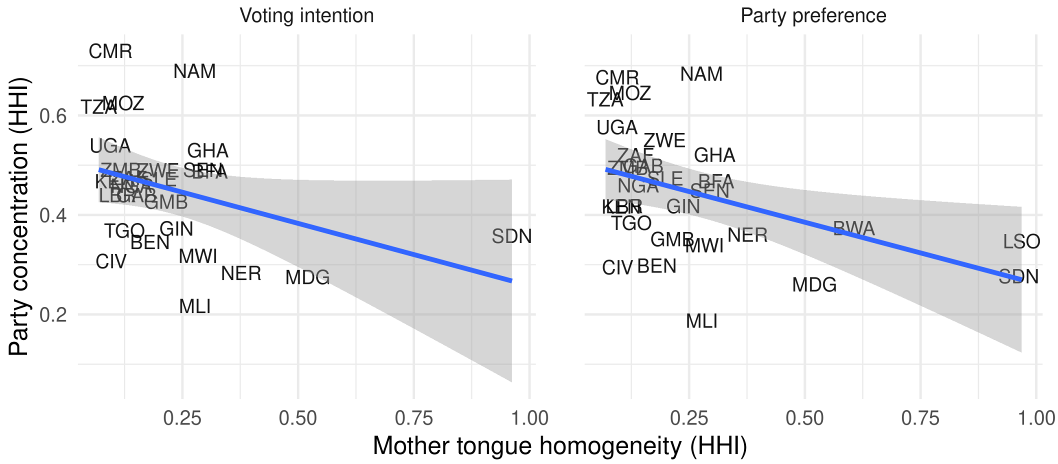



Figure A8: Bi-variate relationship between the national party system concentration and mother tongue homogeneity

Note: Both variables measured as Herfindahl-Hirschman Indices. Based on Afrobarometer Data, round 7.

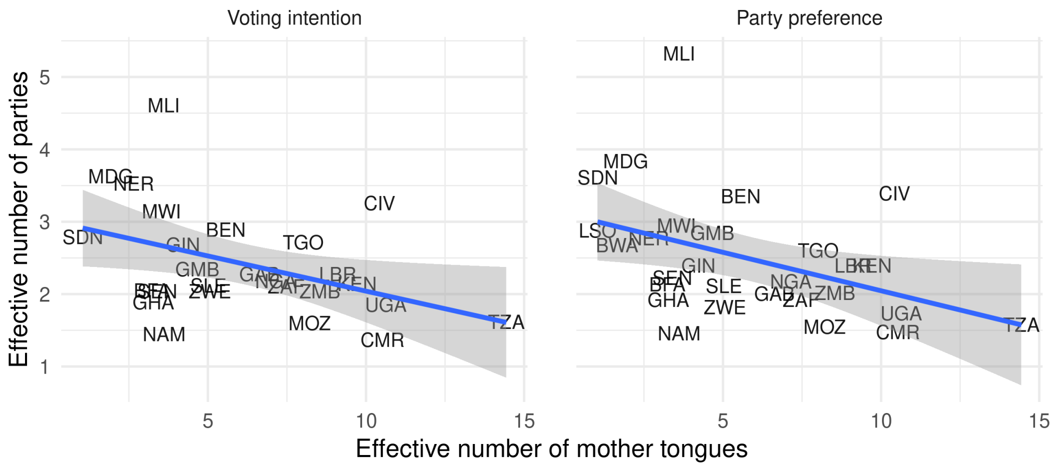



Figure A9: Bi-variate relationship between the national Effective Number of Parties and Effective Number of Mother Tongues

Note: Based on Afrobarometer Data, round 7.

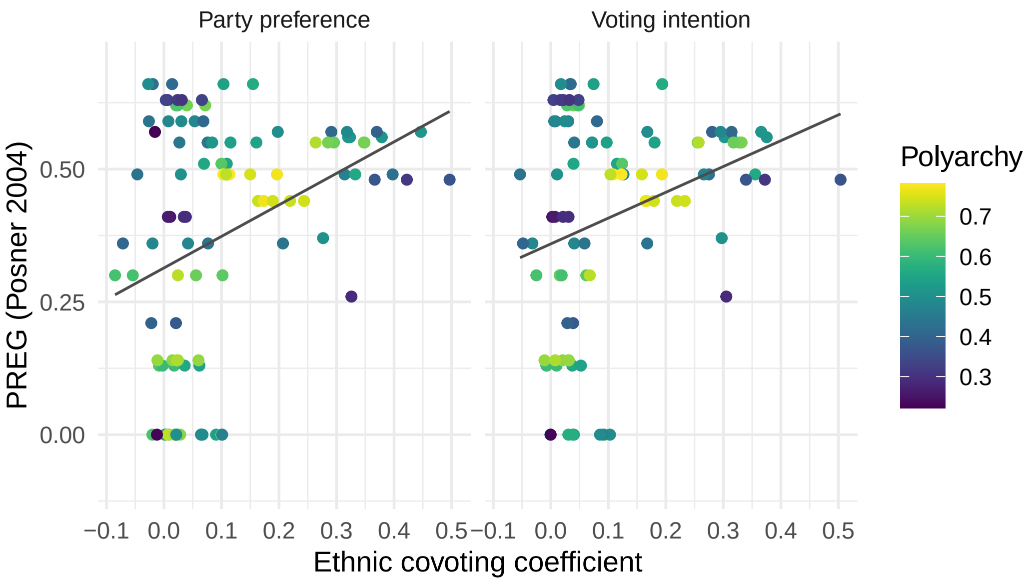



Figure A10: Country-level correlation between effect of shared mother tongues on covoting and PREG Fractionalization Index (Posner 2004).

### D.2 Meso-level: Group- and Party-based measures of ethnic voting

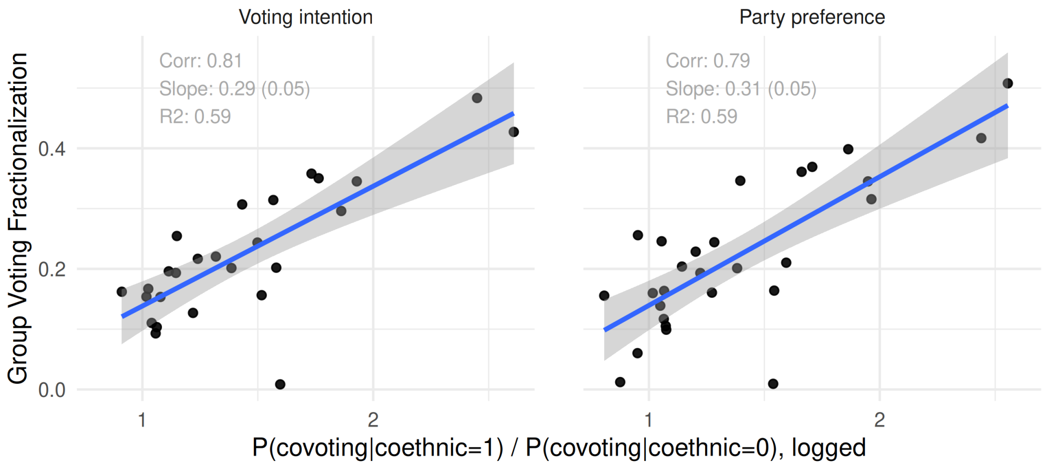



Figure A11: Co-voting among (non-)coethnics and Group Voting Fractionalization (Huber 2012).

Note: Country-level estimates from Afrobarometer Round 7.

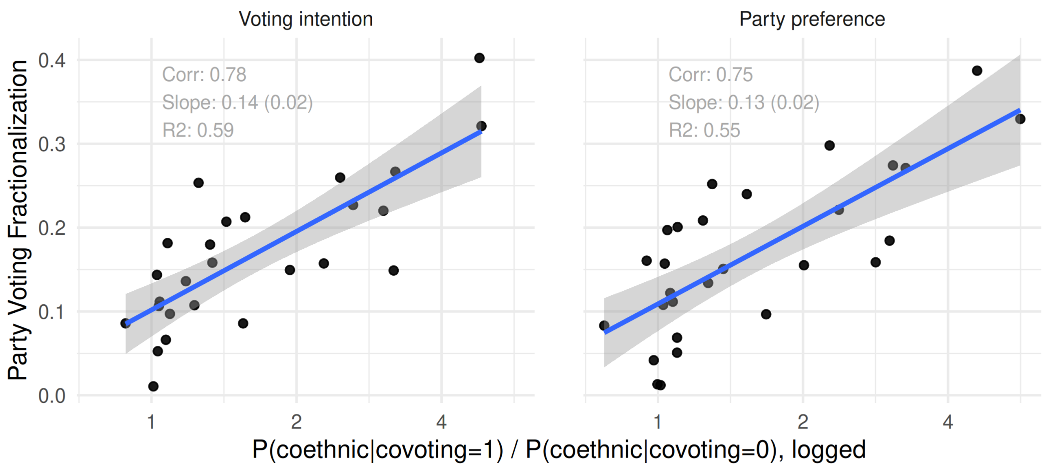



Figure A12: Coethnicity among (non-)covoters and Party Voting Fractionalization (Huber 2012)

Note: Country-level estimates from Afrobarometer Round 7.

## E Heterogeneous effects

### E.1 Micro-level: Ethnic bias, expected favoritism, and political knowledge

We test the moderating effect of individual-level factors on coethnic covoting in a series of interaction models. We measure the prevalence of ethnic bias, by recurring to the Linz-Moreno question (Moreno 1995) which asks whether respondents identify in terms of their national as compared to their ethnic identity on a 1 to 5 scale. Expectations of ethnic favoritism from the government is measured in Afrobarometer Round 5.[16] To test the moderating impact of political knowledge, we construct a knowledge index (0-1) based on the respondents' average knowledge of their MP, local government councilor, the Vice President, the largest party, term limits, and the responsibility for constitutional review as well as a news index based on respondents average frequency of getting news from the radio, TV, newspapers, the internet, and social media (rescaled to 0-1).

Table A4: Covoting intentions and shared mother tongue: Ethnic identification and expected favoritism

| Dependent Variables: | Voting intention | | Party preference | |
|---|---|---|---|---|
| Model: | (1) | (2) | (3) | (4) |
| *Variables* | | | | |
| Shared mother tongue (0/1) | 0.157* | | 0.120+ | |
| | (0.063) | | (0.067) | |
| Ethnic ID (1-5) | 0.009 | | 0.009 | |
| | (0.006) | | (0.007) | |
| Shared mother tongue (0/1) × Ethnic ID (1-5) | 0.002 | | 0.011 | |
| | (0.022) | | (0.021) | |
| Shared ethnicity (0/1) | | 0.133*** | | 0.133*** |
| | | (0.032) | | (0.033) |
| Expect Favors (1-5) | | 0.002 | | -0.001 |
| | | (0.003) | | (0.003) |
| Shared ethnicity (0/1) × Expect Favors (1-5) | | 0.014 | | 0.014 |
| | | (0.009) | | (0.010) |
| *Fixed-effects* | | | | |
| Country x Round | Yes | Yes | Yes | Yes |
| *Fit statistics* | | | | |
| AB Round | 7 | 5 | 7 | 5 |
| Outcome mean | 0.465 | 0.465 | 0.464 | 0.477 |
| Countries | 25 | 25 | 27 | 28 |
| Respondents | 16,752 | 19,250 | 13,155 | 17,544 |
| Dyads | 5,920,273 | 7,700,682 | 3,487,983 | 5,863,955 |
| $R^2$ | 0.087 | 0.087 | 0.086 | 0.103 |
| Within $R^2$ | 0.024 | 0.020 | 0.022 | 0.019 |

*Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1*

The results in Tables A4 and A5 show that ethnic biases, expectations of ethnic

[16]The respective question Q18 asks respondents whether should be "obliged to help their home community or group first" or "should represent everyone [...]," with respondents ranking their agreement with statement 1 vs 2 on a 5-point scale. We recode the variable to capture expected favoritism with higher values.

Table A5: Covoting intentions and shared mother tongue: Political knowledge and news consumption

| Dependent Variables: | Voting intention | | Party preference | |
|---|---|---|---|---|
| Model: | (1) | (2) | (3) | (4) |
| *Variables* | | | | |
| Shared ethnicity (0/1) | 0.131$^{+}$ | | 0.172$^{+}$ | |
| | (0.077) | | (0.090) | |
| Knowledge (0-1) | 0.019 | | 0.071$^{*}$ | |
| | (0.024) | | (0.028) | |
| Shared ethnicity (0/1) $\times$ Knowledge (0-1) | -0.032 | | -0.091 | |
| | (0.076) | | (0.085) | |
| Shared mother tongue (0/1) | | 0.153$^{***}$ | | 0.147$^{**}$ |
| | | (0.044) | | (0.052) |
| News (0-1) | | -0.171$^{***}$ | | -0.147$^{***}$ |
| | | (0.020) | | (0.025) |
| Shared mother tongue (0/1) $\times$ News (0-1) | | 0.047 | | 0.050 |
| | | (0.071) | | (0.097) |
| *Fixed-effects* | | | | |
| Country x Round | Yes | Yes | Yes | Yes |
| *Fit statistics* | | | | |
| AB Round | 3 | 7 | 3 | 7 |
| Outcome mean | 0.514 | 0.460 | 0.532 | 0.458 |
| Countries | 14 | 26 | 16 | 28 |
| Respondents | 10,598 | 17,820 | 10,343 | 13,879 |
| Dyads | 4,134,687 | 6,498,080 | 3,556,471 | 3,810,849 |
| $R^2$ | 0.166 | 0.083 | 0.153 | 0.080 |
| Within $R^2$ | 0.011 | 0.027 | 0.011 | 0.024 |

*Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1*

favoritism, and political knowledge and news consumption do not significantly moderate the effects of coethnicity on shared voting intentions and party preferences.[17] While the effects of ethnic identification and favoritism expectations in Table A4 point in the expected positive direction, the estimates are not statistically significant. The interaction of coethnicity with political knowledge is as expected negative, but again not statistically significant, while higher news consumption (insignificantly) correlates with greater coethnic covoting (Table A4).

### E.2 Meso-level: Groups and parties

Our analysis of meso-level heterogeneity and mechanisms focuses on the importance of politically relevant ethnic group(ing)s as well as differences in covoting

[17] Note that rounds 3 and 5 do not ask for respondents' mother tongue but only their ethnicity.

patterns for parties with and without ethnic constituencies as coded by the VDEM-Party data (Lührmann et al. 2020).

At the group level, we first match the mother tongues of Afrobarometer respondents to politically relevant ethnic groups as listed in the Ethnic Power Relations data (Vogt et al. 2015) as well as Posner's (2004) PREG data. We do so using the LEDA package which makes use of phylogenetic linguistic treeto provide a consistent mapping of ethnic datasets onto each other (Müller-Crepon, Pengl and Bormann 2022).[18] We then encode, for each respondent pair whether the two respondents belong to a common EPR and PREG group, groups which often encompass multiple mother tongues.[19] This coding then allows us to examine by how much the effect of general shared membership in a politically relevant ethnic group(ing) affects covoting above and beyond the effect of a shared mother tongue. Table A6 and A7present the results for the EPR and PREG data respectively. We find in Models 1 and 3 that joint membership in a politically relevant ethnic groups increases covoting by between 16 and 19 percentage points when modelled without accounting for shared mother tongues. When adding the latter indicator in Models 2 and 4, the effect of joint EPR / PREG membership decreases but remains sizeable and significant at between 8 and 10 percentage points. In turn, the effect of shared mother tongues amounts to 10 to 12 percentage points, significantly smaller than in the main specification. Overall, this shows that (1) there are electorally meaningful linguistic cleavages within politically relevant ethnic groups but (2) that these have effects of approximately half the size of the cleavages between politically relevant ethnic groups.[20]

At the level of parties, we test whether the main effect of shared ethnicity on covoting is driven by covoting for parties with or without ethnic constituencies. To that intent, we turn to Afrobarometer Round 6, for which party labels in the responses to the question on party preferences are linked to the VDEM-Party data (Döring and Regel 2019). We then compute, for each party, the share of VDEM-Party experts that state that "the core membership and supporters of this party" belong to one or several "ethnic or racial group(s)" as an indicator of the degree to which a party can be considered an "ethnic party". We cluster parties in three groups, "high" where the majority of coders code it as being ethnic [.5-1], "medium" where a minority does so (.5-0), and "none" where no coder codes it as such. We use this coding to categorize instances of covoting for a given party into the same categories. This coding enables us, in Table A8 to disaggregated the overall effect of coethnicity[21] on covoting (Model 1) into the three categories. The results show that around 60 percent of the overall effect is driven by parties with a high level of ethnic support and approximately 40 percent by parties with a medium level. Importantly, parties coded by no expert has relying on an ethnic support base do not feature higher levels of coethnic covoting.

[18] While EPR and PREG groups are overwhelmingly coarser than Afrobarometer categories, some are not. We then map respondents to all the relevant groups.

[19] We drop pairs where a respondent is not matched to any EPR/PREG group.

[20] For the latter, we have to sum the effect of shared mother tongue and shared EPR / PREG.

[21] Note that round 6 does not ask for respondents' mother tongue but only their ethnicity.

Table A6: Covoting intentions with politically relevant groups: EPR

| Dependent Variables: | Voting intention | | Party preference | |
|---|---|---|---|---|
| Model: | (1) | (2) | (3) | (4) |
| *Variables* | | | | |
| Shared mother tongue (0/1) | | 0.117*** | | 0.105*** |
| | | (0.022) | | (0.023) |
| Shared EPR (0/1) | 0.175*** | 0.100*** | 0.160*** | 0.088*** |
| | (0.039) | (0.028) | (0.041) | (0.025) |
| *Fixed-effects* | | | | |
| Country x Round | Yes | Yes | Yes | Yes |
| *Fit statistics* | | | | |
| Controls | Yes | Yes | Yes | Yes |
| Outcome mean | 0.459 | 0.459 | 0.452 | 0.452 |
| Countries | 17 | 17 | 19 | 19 |
| Respondents | 10,274 | 10,274 | 8,455 | 8,455 |
| Dyads | 3,374,012 | 3,374,012 | 2,109,207 | 2,109,207 |
| $R^2$ | 0.100 | 0.105 | 0.090 | 0.093 |
| Within $R^2$ | 0.032 | 0.038 | 0.024 | 0.028 |

*Standard errors clustered at the level of each respondent's mother tongue in parentheses.*
*Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1*

Table A7: Covoting intentions with politically relevant groups: PREG

| Dependent Variables: | Voting intention | | Party preference | |
|---|---|---|---|---|
| Model: | (1) | (2) | (3) | (4) |
| *Variables* | | | | |
| Shared mother tongue (0/1) | | 0.099*** | | 0.120*** |
| | | (0.022) | | (0.025) |
| Shared PREG (0/1) | 0.185*** | 0.104*** | 0.182*** | 0.083*** |
| | (0.033) | (0.020) | (0.043) | (0.024) |
| *Fixed-effects* | | | | |
| Country x Round | Yes | Yes | Yes | Yes |
| *Fit statistics* | | | | |
| Controls | Yes | Yes | Yes | Yes |
| Outcome mean | 0.466 | 0.466 | 0.453 | 0.453 |
| Countries | 26 | 26 | 28 | 28 |
| Respondents | 11,112 | 11,112 | 9,025 | 9,025 |
| Dyads | 2,915,696 | 2,915,696 | 1,878,800 | 1,878,800 |
| $R^2$ | 0.112 | 0.115 | 0.102 | 0.105 |
| Within $R^2$ | 0.034 | 0.036 | 0.029 | 0.032 |

*Standard errors clustered at the level of each respondent's mother tongue in parentheses.*
*Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1*

Table A8: Covoting by parties' reliance on ethnic support groups

| Dependent Variables: | Party preference | High | Medium | None |
|---|---|---|---|---|
| Model: | (1) | (2) | (3) | (4) |
| *Variables* | | | | |
| Shared ethnicity (0/1) | 0.158*** | 0.096*** | 0.066** | -0.005 |
| | (0.023) | (0.020) | (0.025) | (0.004) |
| *Fixed-effects* | | | | |
| Country x Round | Yes | Yes | Yes | Yes |
| *Fit statistics* | | | | |
| Controls | Yes | Yes | Yes | Yes |
| Outcome mean | 0.503 | 0.179 | 0.222 | 0.103 |
| Dyads | 4,826,832 | 4,826,787 | 4,826,787 | 4,826,787 |
| $R^2$ | 0.104 | 0.403 | 0.412 | 0.368 |
| Within $R^2$ | 0.020 | 0.018 | 0.008 | 0.003 |

*Standard errors clustered at the level of each respondent's ethnicity in parentheses.*
*Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1*

### E.3 Macro-level: Country-level institutions

**Electoral system:** A large literature suggests that proportional electoral systems politicize ethnic identities as particularistic parties face few obstacles to representation and may even join governing coalitions (e.g., Lijphart 2004). In contrast, some majoritarian electoral rules arguably incentivize cross-ethnic mobilization (e.g., Posner 2005). Using data on electoral systems from Bormann and Golder (2022), our results in Table A9 suggest that there are no large or statistically significant difference in the effect of coethnicity on covoting intentions. While PR systems see slightly less shared support for presidential candidates among coethnics at higher average levels of covoting, this difference is not statistically significant. We note that the share of coethnics among covoters (≈25%) as compared to coethnicity among non-covoters (≈17%) is similar in PR and majoritarian systems. This suggests that the (non-)result does not mask differences in the group- and party-based logics of ethnic voting.

**Democracy:** Democratic institutions, specifically competitive elections, are frequently associated with ethnic mobilization (Rabushka and Shepsle 1972; Horowitz 1985; Eifert, Miguel and Posner 2010). As described above, political leaders seek to mobilize majorities through clientelism and patronage, which often follows ethnic lines (Bates 1974), and might reinforce the effect of cultural differences on diverging policy preferences (Lieberman and McClendon 2013). While elections within dictatorships might also follow a clientelist logic, they are less likely to reveal divergent policy preferences. Countering these considerations, we find that the effect of coethnicity on covoting preferences is, if at all, smaller in countries with higher levels of democracy measured via V-DEM's polyarchy index (Coppedge et al. 2016, see Table A10). Again, the differences we observe are not statistically significant. We note that democracies have in general a slightly higher share of coethnic dyads among all dyads than non-democracies. This does however not translate into a statistically different extent to which covoters are more often coethnics than non-coethnics. Again, this suggests that the results reflects an alignment of party- and group-based logics of ethnic voting.

**Traditional institutions** may co-produce local public goods (Baldwin 2016), and act as complements to the state where they are institutionally tied to it (Holzinger et al. 2019; Henn 2022). Therefore, voters have incentives to vote "with their chief" (Baldwin 2013, see also De Kadt and Larreguy 2018). As a result of the entanglement between traditional authorities and ethnic identities, one might expect strong traditional institutions to come with stronger effects of coethnicity on covoting intentions. Using data on the constitutionalization of traditional authorities from (Holzinger et al. 2019), Table A11 shows relatively little and no statistically significant variation in the effect of coethnicity on covoting preferences. As in the previous analyses, further analyses of the group- and party-based interpretations of the CVR suggest that the results reflect both.

Table A9: Covoting and shared mother tongue: By electoral system

| Dependent Variables: | Voting intention | | | Party preference | | |
|---|---|---|---|---|---|---|
| Model: | (1) | (2) | (3) | (4) | (5) | (6) |
| *Variables* | | | | | | |
| Shared mother tongue (0/1) | 0.176*** | 0.124** | 0.207+ | 0.160*** | 0.143* | 0.186* |
| | (0.033) | (0.042) | (0.116) | (0.040) | (0.058) | (0.088) |
| Shared religion (0/1) | 0.030* | -0.001 | 0.007 | 0.027+ | 0.005 | 0.011 |
| | (0.014) | (0.006) | (0.016) | (0.014) | (0.008) | (0.007) |
| Age similarity (decades) | -0.004* | -0.007* | 0.001 | -0.005*** | -0.006 | 0.006+ |
| | (0.002) | (0.003) | (0.001) | (0.001) | (0.004) | (0.004) |
| Shared gender (0/1) | 0.001 | -0.001 | -0.001 | 0.000 | -0.004* | -0.004 |
| | (0.001) | (0.001) | (0.002) | (0.002) | (0.002) | (0.003) |
| Shared education (0/1) | 0.022*** | 0.011 | 0.031* | 0.020*** | -0.008 | 0.023+ |
| | (0.005) | (0.010) | (0.014) | (0.004) | (0.008) | (0.012) |
| Wealth similarity (sd) | 0.006* | 0.004 | -0.002 | 0.008** | -0.001 | -0.008 |
| | (0.002) | (0.006) | (0.008) | (0.003) | (0.004) | (0.006) |
| Shared occupation (0/1) | 0.028*** | 0.031* | 0.023+ | 0.026*** | 0.014 | 0.026+ |
| | (0.008) | (0.015) | (0.012) | (0.007) | (0.009) | (0.013) |
| Geographic proximity (1'000km) | 0.038 | 0.030 | 0.082 | 0.039 | 0.025 | 0.065 |
| | (0.028) | (0.024) | (0.053) | (0.026) | (0.026) | (0.055) |
| Shared urban vs. rural (0/1) | 0.021** | 0.026+ | 0.042*** | 0.020** | 0.018 | 0.035* |
| | (0.007) | (0.014) | (0.007) | (0.007) | (0.012) | (0.015) |
| *Fixed-effects* | | | | | | |
| Country x Round | Yes | Yes | Yes | Yes | Yes | Yes |
| *Fit statistics* | | | | | | |
| Electoral system | FPTP | PR | Mixed | FPTP | PR | Mixed |
| Outcome mean | 0.453 | 0.510 | 0.406 | 0.455 | 0.570 | 0.399 |
| Countries | 15 | 6 | 5 | 16 | 6 | 6 |
| Respondents | 10,903 | 4,113 | 2,816 | 8,634 | 2,280 | 2,974 |
| Dyads | 4,170,640 | 1,482,444 | 853,843 | 2,518,928 | 486,589 | 809,975 |
| $R^2$ | 0.064 | 0.080 | 0.135 | 0.072 | 0.068 | 0.063 |
| Within $R^2$ | 0.023 | 0.015 | 0.051 | 0.020 | 0.018 | 0.032 |

*Standard errors clustered at the level of each respondent's mother tongue in parentheses.*
*Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1*

Table A10: Covoting and shared mother tongue: By level of democracy

| Dependent Variables: | Voting intention | | | Party preference | | |
|---|---|---|---|---|---|---|
| Model: | (1) | (2) | (3) | (4) | (5) | (6) |
| *Variables* | | | | | | |
| Shared mother tongue (0/1) | 0.219** | 0.178** | 0.134*** | 0.214** | 0.190** | 0.110** |
| | (0.073) | (0.053) | (0.029) | (0.068) | (0.059) | (0.039) |
| Shared religion (0/1) | 0.003 | 0.044* | 0.004 | 0.018* | 0.027 | 0.007 |
| | (0.006) | (0.022) | (0.007) | (0.008) | (0.019) | (0.006) |
| Age similarity (decades) | -0.004 | -0.005* | -0.003+ | -0.005* | -0.001 | -0.002 |
| | (0.003) | (0.002) | (0.002) | (0.002) | (0.004) | (0.002) |
| Shared gender (0/1) | 0.002 | -0.003** | 0.001 | 0.000 | -0.002 | 0.001 |
| | (0.001) | (0.001) | (0.001) | (0.001) | (0.002) | (0.003) |
| Shared education (0/1) | 0.027*** | 0.020*** | 0.017* | 0.018** | 0.017* | 0.015+ |
| | (0.008) | (0.006) | (0.008) | (0.006) | (0.008) | (0.008) |
| Wealth similarity (sd) | 0.003 | 0.004 | 0.007 | 0.004 | 0.000 | 0.007 |
| | (0.004) | (0.003) | (0.004) | (0.006) | (0.003) | (0.004) |
| Shared occupation (0/1) | 0.040*** | 0.029** | 0.018 | 0.041*** | 0.021* | 0.013 |
| | (0.009) | (0.009) | (0.011) | (0.011) | (0.008) | (0.009) |
| Geographic proximity (1'000km) | 0.124* | 0.017 | 0.025 | 0.108+ | 0.018 | 0.029 |
| | (0.049) | (0.028) | (0.024) | (0.058) | (0.027) | (0.026) |
| Shared urban vs. rural (0/1) | 0.037* | 0.028*** | 0.012 | 0.047** | 0.020* | 0.005 |
| | (0.015) | (0.005) | (0.009) | (0.015) | (0.008) | (0.006) |
| *Fixed-effects* | | | | | | |
| Country x Round | Yes | Yes | Yes | Yes | Yes | Yes |
| *Fit statistics* | | | | | | |
| Democracy level | Low | Medium | High | Low | Medium | High |
| Outcome mean | 0.469 | 0.436 | 0.475 | 0.480 | 0.432 | 0.471 |
| Countries | 9 | 9 | 8 | 9 | 10 | 9 |
| Respondents | 5,493 | 6,177 | 6,162 | 4,030 | 5,362 | 4,496 |
| Dyads | 1,860,094 | 2,210,861 | 2,435,972 | 1,061,826 | 1,553,381 | 1,200,285 |
| $R^2$ | 0.072 | 0.092 | 0.077 | 0.078 | 0.085 | 0.070 |
| Within $R^2$ | 0.038 | 0.025 | 0.016 | 0.038 | 0.024 | 0.012 |

*Standard errors clustered at the level of each respondent's mother tongue in parentheses.*
*Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1*

Table A11: Covoting and shared mother tongue: By constitutionalization of traditional institutions

| Dependent Variables: | Voting intention | | | Party preference | | |
|---|---|---|---|---|---|---|
| Model: | (1) | (2) | (3) | (4) | (5) | (6) |
| *Variables* | | | | | | |
| Shared mother tongue (0/1) | 0.169* | 0.161*** | 0.181*** | 0.190* | 0.150*** | 0.161* |
| | (0.079) | (0.032) | (0.050) | (0.089) | (0.030) | (0.063) |
| Shared religion (0/1) | 0.033* | 0.031* | -0.008+ | 0.037** | 0.042** | -0.009* |
| | (0.016) | (0.014) | (0.005) | (0.011) | (0.016) | (0.004) |
| Age similarity (decades) | -0.002 | -0.003+ | -0.006* | -0.005* | -0.002 | -0.002 |
| | (0.002) | (0.002) | (0.002) | (0.002) | (0.002) | (0.004) |
| Shared gender (0/1) | -0.003* | 0.000 | 0.002+ | -0.006* | -0.004* | 0.004* |
| | (0.001) | (0.001) | (0.001) | (0.002) | (0.001) | (0.002) |
| Shared education (0/1) | 0.028*** | 0.017* | 0.019*** | 0.033*** | 0.010 | 0.012+ |
| | (0.005) | (0.008) | (0.005) | (0.006) | (0.007) | (0.007) |
| Wealth similarity (sd) | -0.004 | 0.007* | 0.008* | -0.003 | 0.001 | 0.010+ |
| | (0.004) | (0.003) | (0.004) | (0.005) | (0.003) | (0.005) |
| Shared occupation (0/1) | 0.039*** | 0.020+ | 0.029** | 0.045*** | 0.009 | 0.026* |
| | (0.011) | (0.010) | (0.009) | (0.012) | (0.006) | (0.010) |
| Geographic proximity (1'000km) | 0.038 | 0.016 | 0.113** | 0.011 | 0.018 | 0.121** |
| | (0.031) | (0.023) | (0.041) | (0.031) | (0.024) | (0.041) |
| Shared urban vs. rural (0/1) | 0.027*** | 0.018* | 0.034* | 0.027*** | 0.017* | 0.025+ |
| | (0.004) | (0.009) | (0.013) | (0.004) | (0.008) | (0.013) |
| *Fixed-effects* | | | | | | |
| Country x Round | Yes | Yes | Yes | Yes | Yes | Yes |
| *Fit statistics* | | | | | | |
| Trad. Inst. constitut. | Low | Medium | High | Low | Medium | High |
| Outcome mean | 0.395 | 0.470 | 0.508 | 0.426 | 0.447 | 0.486 |
| Countries | 10 | 10 | 6 | 10 | 10 | 8 |
| Respondents | 5,693 | 7,451 | 4,688 | 3,839 | 5,222 | 4,827 |
| Dyads | 1,800,397 | 2,859,008 | 1,847,522 | 874,629 | 1,441,373 | 1,499,490 |
| $R^2$ | 0.095 | 0.087 | 0.033 | 0.123 | 0.078 | 0.049 |
| Within $R^2$ | 0.026 | 0.022 | 0.027 | 0.033 | 0.020 | 0.021 |

*Standard errors clustered at the level of each respondent's mother tongue in parentheses.*
*Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1*

# F Robustness checks

## F.1 Interaction models:

Appendix A.4 shows that the CVR as estimated in the main analysis can potentially suffer from omitted variable bias when significant cross-cleavage interactions are left unmodeled. We address this issue here in two steps. We first model the full set of direct effects of individuals' $i$ and $j$ characteristics and all interactions between them (Eq. A9 on page A5).[22] This amounts to 100s, in some cases more than 1,000 additional parameters[23] which can yield to unstable estimates due to multicollinearity. As a remedy, we follow Beiser-McGrath and Beiser-McGrath (2020) and reduce the set of potential interaction and control variables through a Least Absolute Shrinkage and Selection Operator (LASSO), adding up to a maximum number of 100 additional terms to the main "controls" specification.

We estimate both models for each country covered in round 7 of the Afrobarometer and separately for voting intentions and party preferences. We then compare the results with country-level CVR estimates using the baseline and control specifications. Readers may note that proceeding on a country-by-country basis is a comparatively stringent test, as we introduce significantly more flexibility than in the pooled sample. The approach has the further advantage of facilitating empirical estimation which is challenging as a pooled context comes with many more interaction terms to be estimated on a much larger sample.

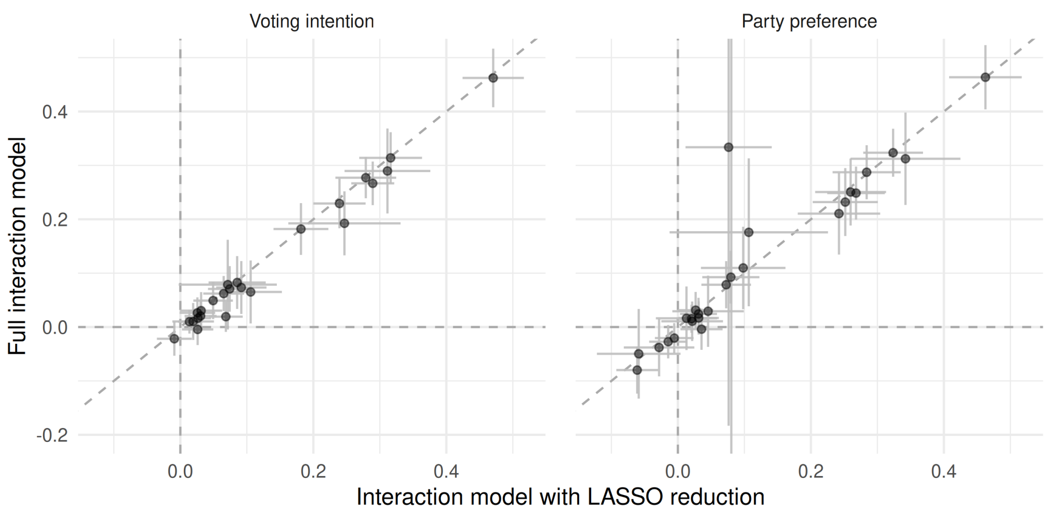



Figure A13: Comparison of country-level effects of coethnicity on covoting from interaction models with the fully set of interactions and LASSO-selected interactions.

Note: Based on Afrobarometer round 7.

[22] We adjust all additional interaction terms to have a mean of zero, such that we can continue to directly interpret the coefficient of coethnicity. To the same intent, interactions between ethnic groups of $i$ and $j$ are shifted such that they have a mean of zero among coethnic pairs if $i$ and $j$ are coethnics and a mean of 0 among non-coethnics if $i$ and $j$ are not coethnic.

[23] This is the case, e.g., in Nigeria where many mother tongues of $i$ interact, e.g., with the many religions and occupations of $j$.

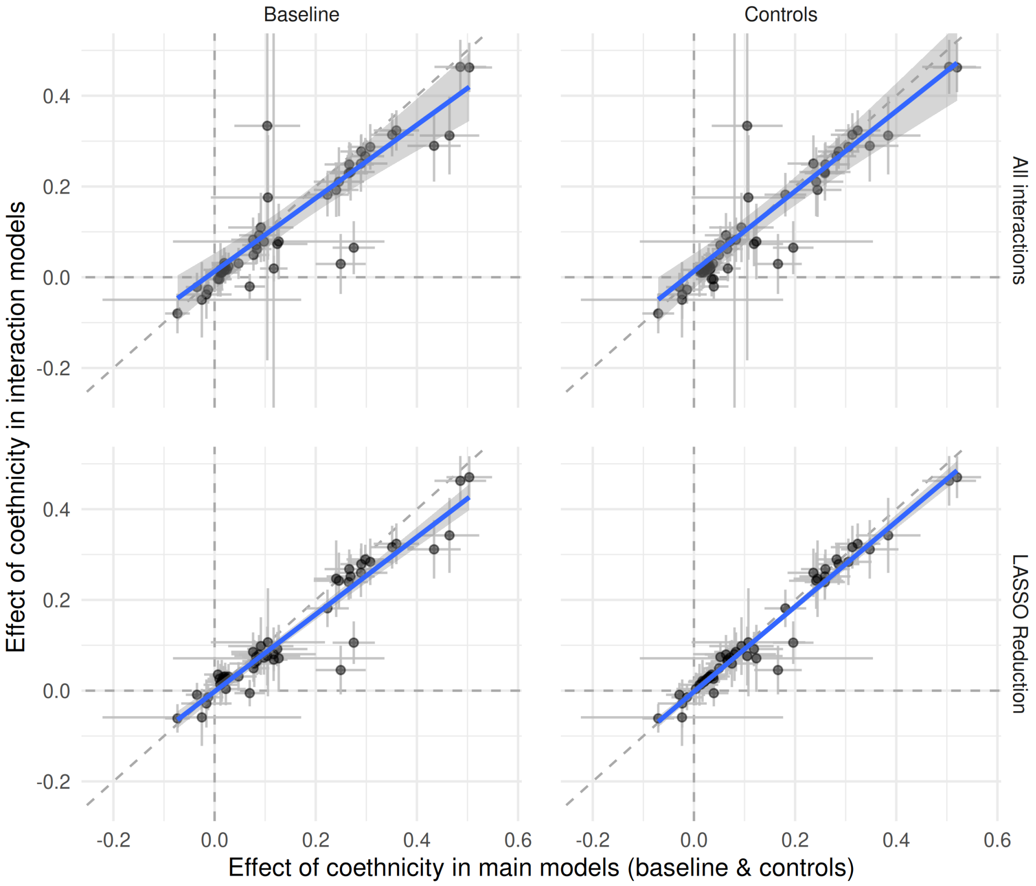



Figure A14: Comparison of country-level effects of coethnicity on covoting from baseline and control specifications and interaction models.

Note: Data from Afrobarometer round 7, separate models for voting intentions and party preference.

We first note the strong overlap between the fully specified and LASSO-reduced models which is shown in Figure A13. In fact, the main divergence consists in a set of one outlier produced through the full interaction model which is due to uncaptured multi-collinearity issues. These results suggest that the LASSO-reduction is successful in streamlining the otherwise overly complicated interaction model.

Figure A14 then plots the results of the interaction models that account for cross-cleavage interactions and our main, country-level results. We first note the close alignment of estimates of the effect of coethnicity with each other as visible in their clustering around the 45-degree line. As to be expected, we observe slightly larger deviations with the empty baseline model than the main "controls" specification. The interaction models yield an effect of coethnicity that amount on average to .85 times the baseline coefficient. In turn, the interaction results are closely aligned with the estimates for the effects of coethnicity using the main controls. Here, estimates from the interaction models on average correspond to .92 times the effects from the models with controls, an estimate which barely significantly different from 1 ($p < .1$).

### F.2 Estimates from multinomial voting models

Based on the discussion in Appendix A.4, we estimate, for each country-round and party $\in P$, a separate OLS model of the probability that survey respondents intend to vote for that party or has a preference for it:[24]

$$P(p_i) = \delta_{0,p} + \theta_p \mathbf{z}_i \tag{A11}$$

where $\delta_0$ denotes a party-specific constant and $\theta_p$ captures the vector of party-specific effects of the covariate vector $\mathbf{z}_i$. $\mathbf{z}_i$ captures individual-level characteristics, in particular $i$ mother-tongue $g \in G$ and control variable $\mathbf{z}'$ used in the main analysis.[25] We use this model to derive for each group, party-specific voting probability $Pr(p|g, \bar{\mathbf{z}}')$ setting all controls $\mathbf{z}'$ to their country-specific mean value. We adjust predicted probabilities to 0 where they are $< 0$ and to 1 where they are $> 1$. Using these predicted voting probabilities, we can derive the effect of coethnicity on covoting as the difference between the expected probabilities that a representative pair of coethnics and non-coethnics covote. Doing so only necessitates weighting group-dyadic covoting probabilities by group's share in the survey population $w_g$:

$$\beta_{coeth} = Pr(p_i = p_j | g_i = g_j, \bar{\mathbf{z}}'_{i,j}) - Pr(p_i = p_j | g_i \neq g_j, \bar{\mathbf{z}}'_{i,j}) \tag{A12}$$

$$= \frac{1}{\sum_g^G w_g^2} \sum_p^P \sum_g^G (w_g^2 \, Pr(p|g, \bar{\mathbf{z}}')^2) - \frac{1}{\sum_{g_i, g_j, g_i \neq g_j}^G w_{g_i} w_{g_j}} \sum_p^P \sum_{g_i, g_j, g_i \neq g_j}^G (w_{g_i} w_{g_j} \, Pr(p|g_i, \bar{\mathbf{z}}') \, Pr(p|g_j, \bar{\mathbf{z}}')) \tag{A13}$$

Essentially, Eq. A13 takes the size-weighted, covoting probability within each ethnic group and substracts from it the size-weighted, pairwise covoting probabilities between all non-identical ethnic groups, all adjusted to the average control values $\bar{\mathbf{z}}$.

This approaches comes with two disadvantages compared to the CVR. First and most importantly, the aggregation of voting probabilities across independently estimated models does not allow for recovering adequate uncertainty estimates. Second, the estimation procedure is comparatively inefficient, as voting probabilities for each party and group have to be computed even though the quantity of interest is based on an average across parties and pairs of groups.

We recover $\beta_{coeth}$ as defined in Eq. A13 for voting intentions and party preferences in each country in round 7 of the Afrobarometer using a baseline model without any control variables as well as the full vector of control defined above. This allows us to directly contrast the results with results from estimating the CVR separately for each country and outcome with the baseline and control specifica-

[24] Note that $P(p_i)$ cal also be estimates using a multinomial logistic regression. Yet, in our case, many ethnic groups are small and do not exhibit positive outcomes for each party, which leads to singularities in the estimation procedure.

[25] Individuals' religious identity (categorical), their age, gender, education level (categorical), wealth index, occupation (categorical), urban status, and latitude and longitude.

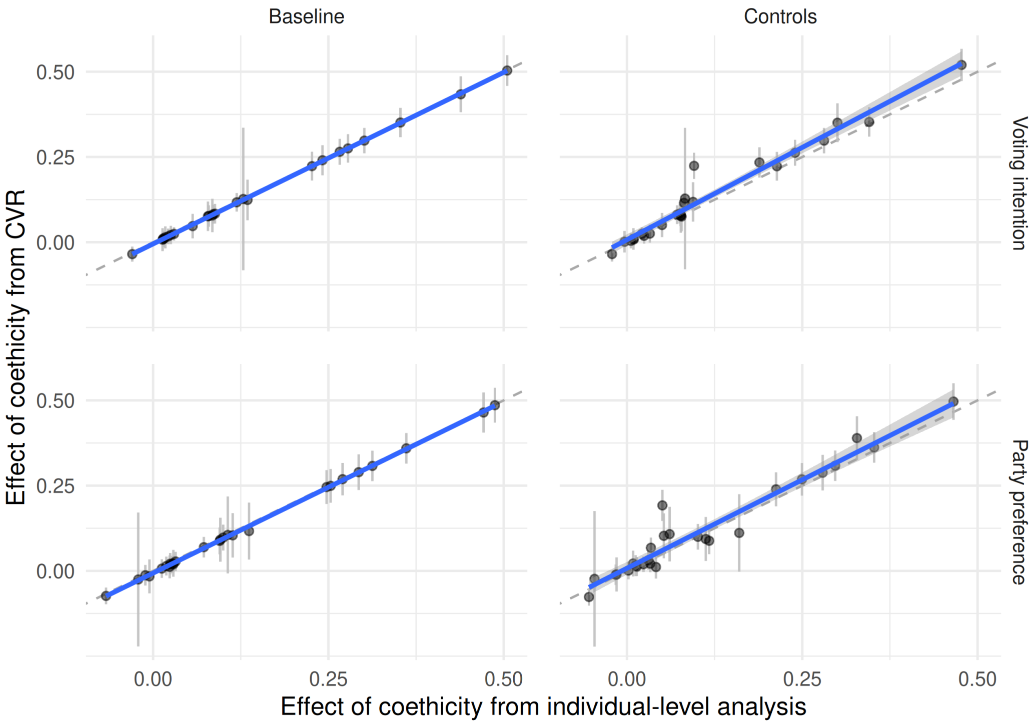



Figure A15: Comparison of country-level effects of coethnicity on covoting from main models compared with estimates recovered from party-by-party "multinomial" OLS estimates of individuals' voting propensities.

Note: Data from Afrobarometer round 7, separate models for voting intentions and party preference. For the individual-level estimation approach, see discussion in text.

tions.

Figure A15 shows the results. Importantly, the baseline specification without controls of the individual-level estimation and the CVR yield almost exactly the same results, thus showing that the aggregation procedure of covariate-adjusted party-level probabilities across groups works as intended. Second, once we include additional control variables the results still coincide substantively with only two country-level estimates deviating significantly from the 45-degree line. Overall, there is a strong linear relationship between the two sets of estimates with only minimal effect overestimation by the CVR. The CVR estimates amount, on average, to 1.06 times the individual level estimate $\beta_{coeth}$ derived above, a deviation is similar to that observed in the fully specified interaction model (Appendix F.1).

### F.3 Accounting for potentially endogenous ethnicity

One threat to inference consists in endogenous ethnic change or identity misreporting among respondents. For example, minority members might be incentivized to report membership in a powerful majority group (Green 2021) or economic incentives might shape political preferences and ethnic groups in parallel (Pengl, Roessler and Rueda 2022). In Appendix F.3, we implement two strategies to gauge in how far such processes can explain our main findings.

First, we leverage differences in the malleability of different ethnic markers. Beyond respondents mother tongue, interviewers in Afrobarometer round 7 also asked respondents about (a) the language spoken in their homes *now* and (b) their "ethnic community, cultural group, or tribe". Both are more malleable than reported mother tongues, with the current language at home being most susceptible to change and strategic reporting. Yet, in particular the ethnicity item is also more precise in reflecting their current ethnic identification than information about individuals' mother tongue, thus reducing measurement error and related downward bias. Re-estimating the CVR using these two variables to construct the indicator of pairwise coethnicity, we find the smallest effect (12 percentage point) for shared language spoken at home. The more precise indicator of shared ethnicity has a slightly larger effect (19 percentage points) than our baseline specification.

Our second strategy draws on the assumption that misreporting and assimilation is least likely to affect pairs of respondents with very distinct and linguistically unrelated mother tongues. We thus estimate the effect of the pairwise linguistic proximity between respondents and find that shared covoting intentions are least likely among respondents who grew up speaking unrelated languages.[26]

As an alternative approach, we vary the coding of coethnicity depending on the level of the language tree on which two mother tongues share the first common node. Very distinct languages come from different languages families (e.g., Khosian or Nilo-Saharan) and share no comon root, whereas more closely related languages – such as Bambara and Fulanke in Malie – share a common node on level 9 of the Mande branch of the Niger-Congo language tree. Recoding coethnicity at the levels of the langauge tree and reestimating the main specification for each recoding shows that we recover positive effects of coethnicity on covoting no matter how coarse our classification of ethnic groups is. While effect estimates decrease with coarser levels (reaching 6 percentage points at level 1), we note that they do so nonlinearly and remain constant above level 5, which already reclassifies many respondent pairs as coethnics which are classified as non-coethnics in the main analysis which is based on the most fine-grained classification of languages. These patterns are well inline with the previous finding that covoting increases with linguistic differences. In combination, these results suggest that strategic misreporting or endogenous ethnic change are unlikely to substantively affect our results.

[26]We compute linguistic distance through the ethnic linkages data from Müller-Crepon, Pengl and Bormann (2022).

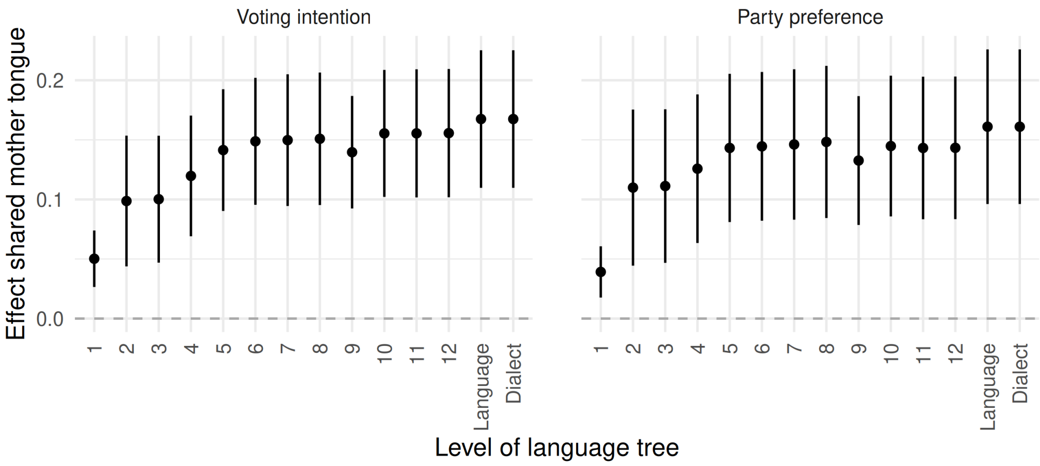



Figure A16: Shared level on language tree, separate models

Table A12: Covoting intentions and shared home language

| Dependent Variables: | Voting intention | | Party preference | |
|---|---|---|---|---|
| Model: | (1) | (2) | (3) | (4) |
| *Variables* | | | | |
| Shared language (0/1) | 0.133*** | 0.124*** | 0.134*** | 0.125*** |
| | (0.029) | (0.028) | (0.030) | (0.029) |
| Shared religion (0/1) | | 0.023** | | 0.022** |
| | | (0.008) | | (0.008) |
| Age similarity (decades) | | -0.004** | | -0.003 |
| | | (0.001) | | (0.002) |
| Shared gender (0/1) | | 0.000 | | -0.001 |
| | | (0.001) | | (0.001) |
| Shared education (0/1) | | 0.020*** | | 0.016** |
| | | (0.005) | | (0.005) |
| Wealth similarity (sd) | | 0.004+ | | 0.003 |
| | | (0.002) | | (0.003) |
| Shared occupation (0/1) | | 0.031*** | | 0.028*** |
| | | (0.008) | | (0.007) |
| Geographic proximity (1'000km) | | 0.051* | | 0.051* |
| | | (0.022) | | (0.021) |
| Shared urban vs. rural (0/1) | | 0.025*** | | 0.024*** |
| | | (0.006) | | (0.006) |
| *Fixed-effects* | | | | |
| Country x Round | Yes | Yes | Yes | Yes |
| *Fit statistics* | | | | |
| Outcome mean | 0.460 | 0.460 | 0.457 | 0.457 |
| Countries | 26 | 26 | 28 | 28 |
| Respondents | 17,716 | 17,716 | 13,799 | 13,799 |
| Dyads | 6,429,441 | 6,429,441 | 3,773,345 | 3,773,345 |
| $R^2$ | 0.070 | 0.073 | 0.070 | 0.073 |
| Within $R^2$ | 0.014 | 0.017 | 0.014 | 0.017 |

*Clustered (lang.round.to & lang.round.from) standard-errors in parentheses*
*Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1*

Table A13: Covoting intentions and shared ethnicity

| Dependent Variables: | Voting intention | | Party preference | |
|---|---|---|---|---|
| Model: | (1) | (2) | (3) | (4) |
| *Variables* | | | | |
| Shared ethnicity (0/1) | 0.201*** | 0.194*** | 0.197*** | 0.190*** |
| | (0.029) | (0.029) | (0.031) | (0.031) |
| Shared religion (0/1) | | 0.017* | | 0.018* |
| | | (0.008) | | (0.008) |
| Age similarity (decades) | | -0.003** | | -0.002+ |
| | | (0.001) | | (0.001) |
| Shared gender (0/1) | | 0.000 | | -0.001 |
| | | (0.001) | | (0.001) |
| Shared education (0/1) | | 0.020*** | | 0.015*** |
| | | (0.004) | | (0.003) |
| Wealth similarity (sd) | | 0.004+ | | 0.003 |
| | | (0.002) | | (0.002) |
| Shared occupation (0/1) | | 0.028*** | | 0.023*** |
| | | (0.006) | | (0.006) |
| Geographic proximity (1'000km) | | 0.030+ | | 0.030+ |
| | | (0.018) | | (0.018) |
| Shared urban vs. rural (0/1) | | 0.025*** | | 0.024*** |
| | | (0.005) | | (0.005) |
| *Fixed-effects* | | | | |
| Country x Round | Yes | Yes | Yes | Yes |
| *Fit statistics* | | | | |
| Outcome mean | 0.459 | 0.459 | 0.457 | 0.457 |
| Countries | 26 | 26 | 28 | 28 |
| Respondents | 17,523 | 17,523 | 13,643 | 13,643 |
| Dyads | 6,301,137 | 6,301,137 | 3,694,387 | 3,694,387 |
| $R^2$ | 0.082 | 0.084 | 0.080 | 0.082 |
| Within $R^2$ | 0.026 | 0.028 | 0.025 | 0.027 |

*Clustered (eth.round.to & eth.round.from) standard-errors in parentheses*
*Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1*

Table A14: Covoting intentions and linguistic proximity

| Dependent Variables: | Voting intention | | Party preference | |
|---|---|---|---|---|
| Model: | (1) | (2) | (3) | (4) |
| *Variables* | | | | |
| Mother tongue proximity (0-1) | 0.193*** | 0.085** | 0.189*** | 0.079* |
| | (0.036) | (0.030) | (0.041) | (0.031) |
| Shared religion (0/1) | 0.024** | 0.020* | 0.024** | 0.020** |
| | (0.008) | (0.008) | (0.007) | (0.008) |
| Age similarity (decades) | -0.004** | -0.004** | -0.003 | -0.003 |
| | (0.001) | (0.001) | (0.002) | (0.002) |
| Shared gender (0/1) | 0.000 | 0.000 | -0.001 | -0.001 |
| | (0.001) | (0.001) | (0.001) | (0.001) |
| Shared education (0/1) | 0.020*** | 0.020*** | 0.016*** | 0.016*** |
| | (0.004) | (0.004) | (0.004) | (0.004) |
| Wealth similarity (sd) | 0.005* | 0.005+ | 0.004 | 0.003 |
| | (0.003) | (0.002) | (0.003) | (0.003) |
| Shared occupation (0/1) | 0.027*** | 0.027*** | 0.024*** | 0.024*** |
| | (0.006) | (0.006) | (0.006) | (0.006) |
| Geographic proximity (1’000km) | 0.043* | 0.031+ | 0.047* | 0.031+ |
| | (0.020) | (0.018) | (0.021) | (0.019) |
| Shared urban vs. rural (0/1) | 0.026*** | 0.026*** | 0.024*** | 0.023*** |
| | (0.005) | (0.006) | (0.006) | (0.006) |
| Shared mother tongue (0/1) | | 0.123*** | | 0.122*** |
| | | (0.028) | | (0.031) |
| *Fixed-effects* | | | | |
| Country x Round | Yes | Yes | Yes | Yes |
| *Fit statistics* | | | | |
| Outcome mean | 0.459 | 0.459 | 0.457 | 0.457 |
| Countries | 26 | 26 | 28 | 28 |
| Respondents | 17,696 | 17,696 | 13,793 | 13,793 |
| Dyads | 6,421,650 | 6,421,650 | 3,773,003 | 3,773,003 |
| $R^2$ | 0.078 | 0.082 | 0.075 | 0.079 |
| Within $R^2$ | 0.021 | 0.026 | 0.019 | 0.023 |

*Standard errors clustered at the level of each respondent’s mother tongue in parentheses.*
*Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1*

## F.4 Accounting for geographic variation

A second threat to inference originates in the geography of ethnic groups. Many ethnic groups' inhabit regionally distinct homelands. Coethnic voting intentions might simply emerge from an alignment of political preferences of individuals who live in the same administrative region or even location based on joint economic interests (Boone et al. 2022; Boone 2024), an argument that dovetails with findings of non-ethnic voting of local minorities in presidential elections (Ichino and Nathan 2013). This risk is further compounded by previous findings that the drawing of subnational borders has partially shaped ethnic geography itself (Posner 2005; Müller-Crepon 2023). Yet, we note that ethnic geography likewise affects individuals' migration decisions and place of residence (Müller-Crepon 2023, see also Marbach 2021) as well as countries' division into administrative units (Muller-Crepon 2024).

We address this threat in two steps. First, Table A15 assess whether controlling for more finegrained geographical commonalities affects the estimated effect of coethnicity and whether the geography and coethnicity interact. Models 1 and 4 show, for voting intentions and party preferences, respectively, that geographic proximity has a non-linear association with covoting, with covoting becoming increasingly less common the more distant to respondents are. Accounting for this pattern does however not substantively change the effect of coethnicity. Models 2 and 5 then show that adding dummy variables for respondents who live in the same enumeration area, district, and region does not impact the estimate of coethnicity either, even though individuals from the same areas tends to covote more often. Lastly, we do not find substantively large or consistent interaction effects between joint residence patterns and coethnicity. If at all, the effect of coethnicity is reduced by about 4 percentage points within the same enumeration area,[27] yet this pattern does not replicate for the district or regional level. Overall then, accounting for additional nuances of shared geographic backgrounds between respondents does not explain our results.

In a second step, we excluding variation between administrative regions or single localities from our data altogether. We do so by constructing our dyadic comparisons *after* splitting each country-round into disjoint samples from (a) administrative regions and (b) enumeration areas (EAs).[28] The resulting data then features no dyads that span across these spatial units, leaving only comparisons among respondents who live in the same region/EA. Doing so increases the rate of shared mother tongues from 22 percent in the full sample to 42 percent within regions and 65 percent within enumeration areas.[29] Equally important, we note that *non-coethnic* dyads within regions and EAs are not representative of the overall distribution of respondent pairs. In particular, they feature greater linguistic proximity (average distance of .2 compared to .5), which should increase covoting among non-coethnics and water down the effect of coethnicity.

[27] This coincides with earlier findings by Ichino and Nathan (2013) and findings below.

[28] For countries for which we subsample large Afrobarometer surveys to 1'200 respondents, we repeat this procedure, but sampling enumeration areas in a block-wise manner to maintain as large a sample as possible. Typically, there are 8 respondents per EA.

[29] Similarly, covoting intentions increase from 46 percent to 51 and 59 percent within regions and enumeration areas, respectively.

Table A15: Co-voting intentions and shared mother tongue: Distance and administrative geography

| Dependent Variables: | Voting intention | | | Party preference | | |
|---|---|---|---|---|---|---|
| Model: | (1) | (2) | (3) | (4) | (5) | (6) |
| *Variables* | | | | | | |
| Shared mother tongue (0/1) | 0.164*** | 0.164*** | 0.163*** | 0.160*** | 0.159*** | 0.160*** |
| | (0.030) | (0.029) | (0.030) | (0.033) | (0.033) | (0.035) |
| Geographic proximity (1'000km) | -0.155* | 0.051* | 0.051* | -0.130+ | 0.033 | 0.032 |
| | (0.063) | (0.023) | (0.023) | (0.070) | (0.024) | (0.024) |
| Geographic proximity (1'000km) square | 0.058** | | | 0.049* | | |
| | (0.021) | | | (0.021) | | |
| Shared enum. area | | 0.044** | 0.067*** | | 0.028 | 0.051*** |
| | | (0.015) | (0.010) | | (0.017) | (0.012) |
| Shared district (adm2) | | -0.036** | -0.045*** | | -0.024+ | -0.030** |
| | | (0.013) | (0.009) | | (0.014) | (0.009) |
| Shared region (adm1) | | 0.034* | 0.035* | | 0.048** | 0.052** |
| | | (0.015) | (0.015) | | (0.017) | (0.018) |
| Shared mother tongue (0/1) × Shared enum. area | | | -0.043* | | | -0.036+ |
| | | | (0.017) | | | (0.021) |
| Shared mother tongue (0/1) × Shared district (adm2) | | | 0.022 | | | 0.014 |
| | | | (0.015) | | | (0.017) |
| Shared mother tongue (0/1) × Shared region (adm1) | | | -0.002 | | | -0.011 |
| | | | (0.016) | | | (0.018) |
| *Fit statistics* | | | | | | |
| Country-Round FE | Yes | Yes | Yes | Yes | Yes | Yes |
| Controls | Yes | Yes | Yes | Yes | Yes | Yes |
| Outcome mean | 0.460 | 0.460 | 0.460 | 0.458 | 0.465 | 0.465 |
| Countries | 26 | 26 | 26 | 28 | 27 | 27 |
| Respondents | 17,832 | 17,760 | 17,760 | 13,888 | 13,218 | 13,218 |
| Dyads | 6,506,927 | 6,450,456 | 6,450,456 | 3,815,492 | 3,596,019 | 3,596,019 |
| $R^2$ | 0.080 | 0.081 | 0.081 | 0.078 | 0.080 | 0.080 |
| Within $R^2$ | 0.024 | 0.024 | 0.024 | 0.022 | 0.024 | 0.024 |

*Standard errors clustered at the level of each respondent's mother tongue in parentheses.*
*Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1*

While decreasing in size, the estimated effects of a shared mother tongue remain sizeable even within regions (12-14ppts) and enumeration areas (10-12ppts) when using country-round fixed effects. Once we add region and EA fixed effects, estimates decrease by another 2-4 percentage points to 10-11 percentage pints within regions and 7-8 percentage points within EAs. This decrease from the main effect suggests more frequent alignment of electoral preferences across ethnic lines within small geographic radii within which coethnicity and covoting are both generally more common. Yet, we note that respondent pairs within EAs (regions) only make up a small fraction of around 6 (15) percent of the full data, thus having only relatively minor impact on the overall results.

Figure A17 investigates whether the drop in the estimated effect of shared mother tongues varies between countries. The results show that country-level estimates are quite closely aligned across specifications. Similar to the results shown in Table A17, the EA-level specifications lead to a drop in effect sizes by 34-40 percent when using country FEs and by 52-55 percent when using EA FEs. While in states such as Kenya the EA-specific effect of coethnicity is almost as large as that in the overall sample, Zambia for example sees a drop from around .2 to 0, but only once we include EA FEs wich soak up most of the variation in covoting, which stands at an average of 70 percent within EAs but 35 percent nationwide.[30]

[30] The difference between the rate of covoting within EAs and nationwide is indeed strongly related to the difference in the estimated effect of coethnicity ($\beta = .43, p < .001$).

Table A16: Covoting intentions and shared mother tongue: Within Administrative Regions

| Dependent Variables: | Voting intention | | Party preference | |
|---|---|---|---|---|
| Model: | (1) | (2) | (3) | (4) |
| *Variables* | | | | |
| Shared mother tongue (0/1) | 0.137*** | 0.107*** | 0.117*** | 0.101*** |
| | (0.020) | (0.018) | (0.023) | (0.019) |
| Shared religion (0/1) | 0.020* | 0.017** | 0.024* | 0.018* |
| | (0.008) | (0.006) | (0.011) | (0.008) |
| Age similarity (decades) | -0.007*** | -0.005*** | -0.007** | -0.005* |
| | (0.002) | (0.001) | (0.003) | (0.002) |
| Shared gender (0/1) | 0.001 | 0.002 | 0.000 | 0.001 |
| | (0.002) | (0.002) | (0.002) | (0.002) |
| Shared education (0/1) | 0.026*** | 0.022*** | 0.018*** | 0.012** |
| | (0.006) | (0.005) | (0.005) | (0.004) |
| Wealth similarity (sd) | 0.009* | 0.008+ | 0.006 | 0.006 |
| | (0.004) | (0.004) | (0.004) | (0.004) |
| Shared occupation (0/1) | 0.021* | 0.010 | 0.018* | 0.014+ |
| | (0.009) | (0.008) | (0.009) | (0.008) |
| Geographic proximity (1'000km) | 0.007 | 0.122 | 0.067 | 0.140 |
| | (0.065) | (0.097) | (0.070) | (0.093) |
| Shared urban vs. rural (0/1) | 0.024*** | 0.030*** | 0.025* | 0.031*** |
| | (0.007) | (0.006) | (0.010) | (0.008) |
| *Fixed-effects* | | | | |
| Country x Round | Yes | | Yes | |
| Region | | Yes | | Yes |
| *Fit statistics* | | | | |
| Outcome mean | 0.518 | 0.518 | 0.521 | 0.521 |
| Regions | 324 | 324 | 359 | 359 |
| Respondents | 17,831 | 17,831 | 13,887 | 13,887 |
| Dyads | 1,008,016 | 1,008,016 | 600,735 | 600,735 |
| $R^2$ | 0.080 | 0.135 | 0.081 | 0.139 |
| Within $R^2$ | 0.021 | 0.014 | 0.016 | 0.012 |

*Standard errors clustered at the level of each respondent's mother tongue in parentheses.*
*Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1*

Table A17: Covoting intentions and shared mother tongue: Within Enumeration Areas

| Dependent Variables: | Voting intention | | Party preference | |
|---|---|---|---|---|
| Model: | (1) | (2) | (3) | (4) |
| *Variables* | | | | |
| Shared mother tongue (0/1) | 0.111*** | 0.067*** | 0.096*** | 0.076*** |
| | (0.016) | (0.017) | (0.015) | (0.018) |
| Shared religion (0/1) | 0.031** | 0.028*** | 0.028* | 0.019+ |
| | (0.011) | (0.008) | (0.012) | (0.010) |
| Age similarity (decades) | -0.009*** | -0.002 | -0.003 | -0.001 |
| | (0.003) | (0.002) | (0.003) | (0.003) |
| Shared gender (0/1) | 0.003 | 0.001 | -0.010+ | -0.010+ |
| | (0.005) | (0.004) | (0.006) | (0.005) |
| Shared education (0/1) | 0.032*** | 0.020*** | 0.025** | 0.014+ |
| | (0.007) | (0.006) | (0.009) | (0.008) |
| Wealth similarity (sd) | 0.016** | 0.015** | 0.009 | 0.009 |
| | (0.005) | (0.005) | (0.006) | (0.006) |
| Shared occupation (0/1) | 0.041*** | 0.004 | 0.049*** | 0.015* |
| | (0.008) | (0.006) | (0.010) | (0.007) |
| *Fixed-effects* | | | | |
| Country x Round | Yes | | Yes | |
| Enum. Area | | Yes | | Yes |
| *Fit statistics* | | | | |
| Outcome mean | 0.596 | 0.596 | 0.590 | 0.590 |
| Enum. areas | 3,659 | 3,659 | 3,372 | 3,372 |
| Respondents | 17,584 | 17,584 | 13,224 | 13,224 |
| Dyads | 39,259 | 39,259 | 23,398 | 23,398 |
| $R^2$ | 0.073 | 0.364 | 0.081 | 0.418 |
| Within $R^2$ | 0.017 | 0.005 | 0.013 | 0.005 |

*Standard errors clustered at the level of each respondent's mother tongue in parentheses.*
*Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1*

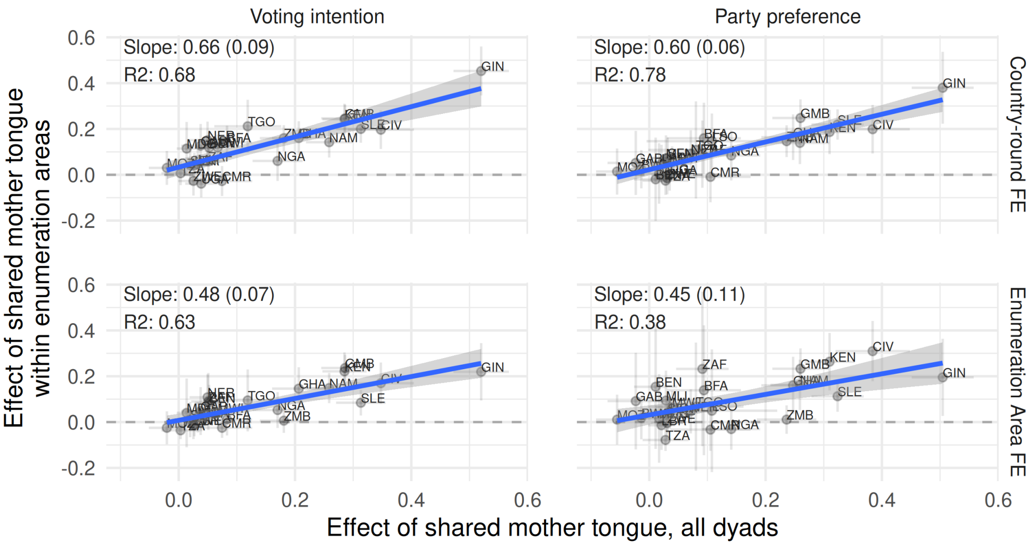



Figure A17: Country-wise estimate of effect of shared mother tongue on covoting across the whole sample (x-axis) and within enumeration areas only (y-axis)

Note: Coefficients result from the fully specified model in Equation 10 estimated by country. Specifications within Enumeration Areas correspond to models in Table A17. We drop Sudan which features an outlier estimates of .4 in the EA-level analysis due to low variation in coethnicity in the mostly Arabic-speaking country.

### F.5 Expanding the economic wealth index

Table A18: Co-voting and shared mother tongue

| Dependent Variables: | Voting intention | Party preference |
|---|---|---|
| Model: | (1) | (2) |
| *Variables* | | |
| Shared mother tongue (0/1) | 0.168*** | 0.163*** |
| | (0.029) | (0.033) |
| Shared religion (0/1) | 0.020* | 0.019** |
| | (0.008) | (0.007) |
| Age similarity (decades) | -0.003* | -0.002 |
| | (0.001) | (0.002) |
| Shared gender (0/1) | 0.000 | -0.001 |
| | (0.001) | (0.001) |
| Shared education (0/1) | 0.020*** | 0.016*** |
| | (0.004) | (0.003) |
| Class index similarity (sd) | 0.002 | 0.002 |
| | (0.002) | (0.002) |
| Shared occupation (0/1) | 0.028*** | 0.024*** |
| | (0.006) | (0.006) |
| Geographic proximity (1'000km) | 0.040* | 0.036+ |
| | (0.020) | (0.019) |
| Shared urban vs. rural (0/1) | 0.024*** | 0.022*** |
| | (0.005) | (0.006) |
| *Fixed-effects* | | |
| Country x Round | Yes | Yes |
| *Fit statistics* | | |
| Outcome mean | 0.458 | 0.454 |
| Countries | 26 | 28 |
| Respondents | 17,806 | 13,837 |
| Dyads | 6,484,734 | 3,782,804 |
| $R^2$ | 0.079 | 0.076 |
| Within $R^2$ | 0.023 | 0.021 |

*Standard errors clustered at the level of each respondent's mother tongue in parentheses.*
*Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1*

## F.6 Data construction

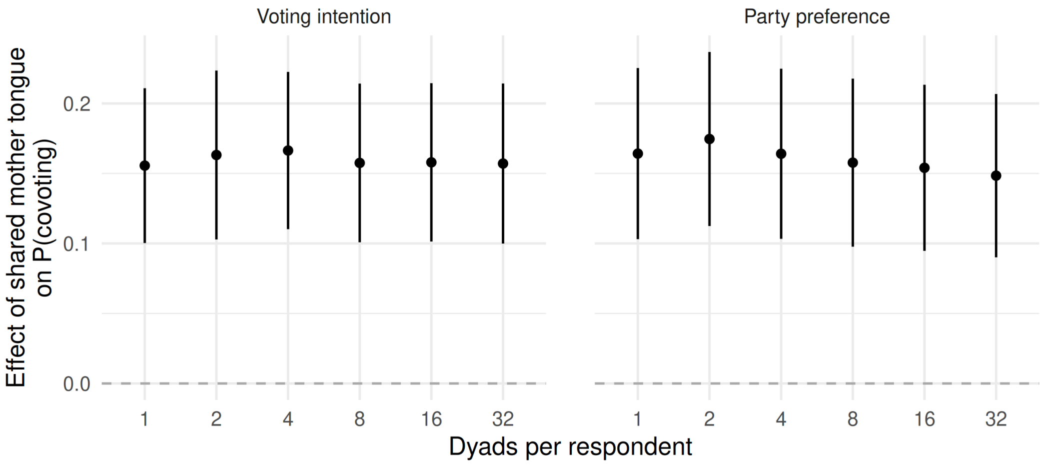



Figure A18: Effect of shared mother tongue by number of comparisons per respondent

Note: Coefficients result from the fully specified model in Equation 10 estimated with different number of dyadic pairings for each respondents, and using respondents' self-identified ethnicity to construct the co-ethnicity indicator. The estimated are thus equivalent to Models 2 and 4 in Table 1 in all respects except for the construction of dyads.

Table A19: Covoting intentions and shared mother tongue: Country-weights

| Dependent Variables: | Voting intention | | Party preference | |
|---|---|---|---|---|
| Model: | (1) | (2) | (3) | (4) |
| *Variables* | | | | |
| Shared mother tongue (0/1) | 0.190*** | 0.182*** | 0.192*** | 0.185*** |
| | (0.031) | (0.031) | (0.041) | (0.042) |
| Shared religion (0/1) | | 0.015* | | 0.010 |
| | | (0.006) | | (0.007) |
| Age similarity (decades) | | -0.003** | | -0.003 |
| | | (0.001) | | (0.002) |
| Shared gender (0/1) | | 0.001 | | 0.000 |
| | | (0.001) | | (0.001) |
| Shared education (0/1) | | 0.021*** | | 0.018*** |
| | | (0.004) | | (0.004) |
| Wealth similarity (sd) | | 0.005* | | 0.004 |
| | | (0.002) | | (0.003) |
| Shared occupation (0/1) | | 0.027*** | | 0.026*** |
| | | (0.006) | | (0.007) |
| Geographic proximity (1'000km) | | 0.037+ | | 0.041+ |
| | | (0.021) | | (0.021) |
| Shared urban vs. rural (0/1) | | 0.023*** | | 0.024*** |
| | | (0.006) | | (0.007) |
| *Fixed-effects* | | | | |
| Country x Round | Yes | Yes | Yes | Yes |
| *Fit statistics* | | | | |
| Outcome mean | 0.460 | 0.460 | 0.458 | 0.458 |
| Countries | 26 | 26 | 28 | 28 |
| Respondents | 17,832 | 17,832 | 13,888 | 13,888 |
| Dyads | 6,506,927 | 6,506,927 | 3,815,492 | 3,815,492 |
| $R^2$ | 0.077 | 0.079 | 0.075 | 0.077 |
| Within $R^2$ | 0.024 | 0.026 | 0.024 | 0.027 |

*Standard errors clustered at the level of each respondent's mother tongue in parentheses.*
*Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1*

Table A20: Covoting intentions and shared mother tongue: Recoding 'other' parties as single parties

| Dependent Variables: | Voting intention | | Party preference | |
|---|---|---|---|---|
| Model: | (1) | (2) | (3) | (4) |
| *Variables* | | | | |
| Shared mother tongue (0/1) | 0.163*** | 0.156*** | 0.168*** | 0.160*** |
| | (0.029) | (0.029) | (0.033) | (0.033) |
| Shared religion (0/1) | | 0.019* | | 0.021** |
| | | (0.007) | | (0.008) |
| Age similarity (decades) | | -0.003* | | -0.003 |
| | | (0.001) | | (0.002) |
| Shared gender (0/1) | | 0.000 | | -0.002 |
| | | (0.001) | | (0.001) |
| Shared education (0/1) | | 0.020*** | | 0.017*** |
| | | (0.004) | | (0.004) |
| Wealth similarity (sd) | | 0.004 | | 0.003 |
| | | (0.002) | | (0.003) |
| Shared occupation (0/1) | | 0.027*** | | 0.026*** |
| | | (0.006) | | (0.006) |
| Geographic proximity (1'000km) | | 0.031+ | | 0.037+ |
| | | (0.018) | | (0.019) |
| Shared urban vs. rural (0/1) | | 0.026*** | | 0.023*** |
| | | (0.006) | | (0.006) |
| *Fixed-effects* | | | | |
| Country x Round | Yes | Yes | Yes | Yes |
| *Fit statistics* | | | | |
| Outcome mean | 0.423 | 0.423 | 0.446 | 0.446 |
| Countries | 28 | 28 | 28 | 28 |
| Respondents | 18,714 | 18,714 | 14,128 | 14,128 |
| Dyads | 7,074,839 | 7,074,839 | 3,914,781 | 3,914,781 |
| $R^2$ | 0.086 | 0.089 | 0.079 | 0.082 |
| Within $R^2$ | 0.018 | 0.021 | 0.019 | 0.021 |

*Standard errors clustered at the level of each respondent's mother tongue in parentheses.*
*Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1*

## F.7 Model specification

Table A21: Covoting intentions and shared mother tongue: Logistic regression

| Dependent Variables: | Voting intention | | Party preference | |
|---|---|---|---|---|
| Model: | (1) | (2) | (3) | (4) |
| *Variables* | | | | |
| Shared mother tongue (0/1) | 0.772*** | 0.736*** | 0.745*** | 0.711*** |
| | (0.132) | (0.131) | (0.147) | (0.146) |
| Shared religion (0/1) | | 0.086* | | 0.087* |
| | | (0.036) | | (0.034) |
| Age similarity (decades) | | -0.017** | | -0.011 |
| | | (0.005) | | (0.008) |
| Shared gender (0/1) | | 0.000 | | -0.004 |
| | | (0.003) | | (0.005) |
| Shared education (0/1) | | 0.091*** | | 0.073*** |
| | | (0.017) | | (0.016) |
| Wealth similarity (sd) | | 0.020+ | | 0.014 |
| | | (0.010) | | (0.011) |
| Shared occupation (0/1) | | 0.124*** | | 0.109*** |
| | | (0.025) | | (0.025) |
| Geographic proximity (1'000km) | | 0.176* | | 0.176* |
| | | (0.086) | | (0.085) |
| Shared urban vs. rural (0/1) | | 0.109*** | | 0.099*** |
| | | (0.024) | | (0.026) |
| *Fixed-effects* | | | | |
| Country x Round | Yes | Yes | Yes | Yes |
| *Fit statistics* | | | | |
| Outcome mean | 0.460 | 0.460 | 0.458 | 0.458 |
| Countries | 26 | 26 | 28 | 28 |
| Respondents | 17,832 | 17,832 | 13,888 | 13,888 |
| Dyads | 6,506,927 | 6,506,927 | 3,815,492 | 3,815,492 |

*Standard errors clustered at the level of each respondent's mother tongue in parentheses.*
*Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1*

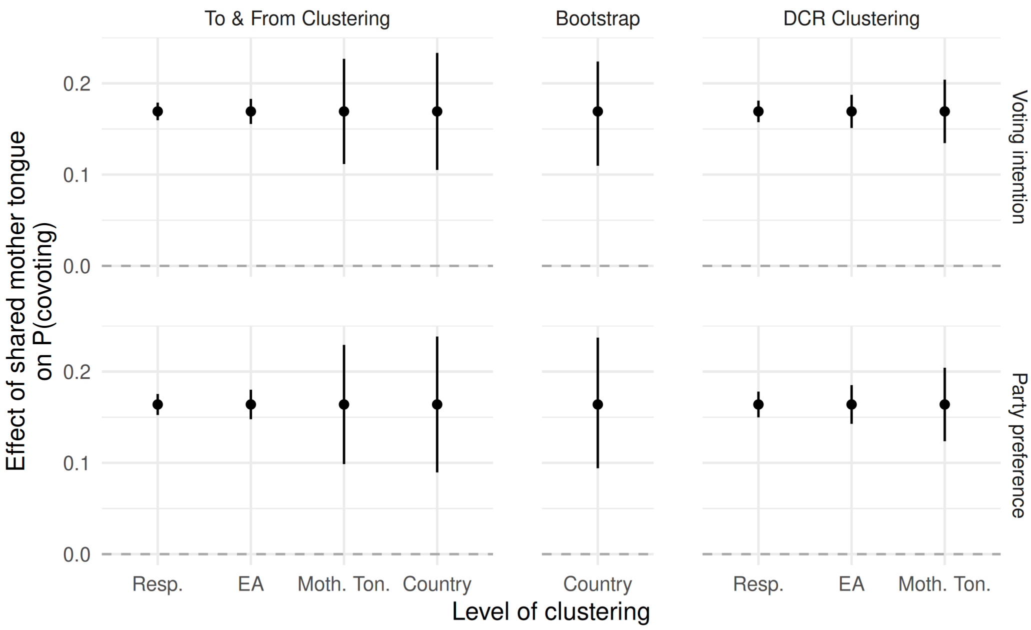



Figure A19: Varying the clustering of standard errors

Note: Coefficients result from the fully specified model in Equation 10 estimated with different standard errors, and using respondents' self-identified ethnicity to construct the co-ethnicity indicator. The estimated models are thus equivalent to Models 2 and 4 in Table 1 in all respects except for the standard errors.

Table A22: Covoting intentions and shared mother tongue: Fixed effects specification

| Dependent Variables: | Voting intention | | | Party preference | | |
|---|---|---|---|---|---|---|
| Model: | (1) | (2) | (3) | (4) | (5) | (6) |
| *Variables* | | | | | | |
| Shared mother tongue (0/1) | 0.157*** | 0.152*** | 0.152*** | 0.174*** | 0.159*** | 0.171*** |
| | (0.032) | (0.032) | (0.033) | (0.036) | (0.034) | (0.037) |
| Shared religion (0/1) | 0.017* | 0.021** | 0.029** | 0.016* | 0.019* | 0.025** |
| | (0.007) | (0.008) | (0.009) | (0.008) | (0.008) | (0.009) |
| Age similarity (decades) | -0.004*** | -0.002* | 0.003*** | -0.003 | -0.001 | 0.002*** |
| | (0.001) | (0.001) | (0.001) | (0.002) | (0.001) | (0.001) |
| Shared gender (0/1) | 0.000 | 0.001+ | 0.002** | -0.001 | 0.001 | 0.003** |
| | (0.001) | (0.001) | (0.001) | (0.001) | (0.001) | (0.001) |
| Shared education (0/1) | 0.019*** | 0.015*** | 0.009*** | 0.014*** | 0.011*** | 0.008*** |
| | (0.004) | (0.003) | (0.002) | (0.004) | (0.003) | (0.002) |
| Wealth similarity (sd) | 0.004* | 0.005*** | 0.005*** | 0.004+ | 0.006*** | 0.006*** |
| | (0.002) | (0.001) | (0.001) | (0.002) | (0.002) | (0.001) |
| Shared occupation (0/1) | 0.026*** | 0.007* | 0.003* | 0.025*** | 0.009*** | 0.003* |
| | (0.006) | (0.003) | (0.001) | (0.005) | (0.002) | (0.001) |
| Geographic proximity (1’000km) | 0.062** | 0.122*** | 0.121*** | 0.066** | 0.118*** | 0.111*** |
| | (0.022) | (0.032) | (0.030) | (0.022) | (0.030) | (0.029) |
| Shared urban vs. rural (0/1) | 0.020*** | 0.002 | 0.003 | 0.018*** | 0.004 | 0.004+ |
| | (0.005) | (0.002) | (0.002) | (0.005) | (0.002) | (0.002) |
| *Fit statistics* | | | | | | |
| Fixed Effects | Lang. | EA | Resp. | Lang. | EA | Resp. |
| Outcome mean | 0.460 | 0.460 | 0.460 | 0.458 | 0.458 | 0.458 |
| Countries | 26 | 26 | 26 | 28 | 28 | 28 |
| Respondents | 17,832 | 17,832 | 17,832 | 13,888 | 13,888 | 13,888 |
| Dyads | 6,506,927 | 6,506,927 | 6,506,927 | 3,815,492 | 3,815,492 | 3,815,492 |
| $R^2$ | 0.122 | 0.225 | 0.426 | 0.124 | 0.246 | 0.436 |
| Within $R^2$ | 0.017 | 0.021 | 0.026 | 0.019 | 0.022 | 0.029 |

*Standard errors clustered at the level of each respondent’s mother tongue in parentheses.*
*Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1*

### F.8 Network-based partition model

We here apply the recently developed *Probabilistic Spatial Partition Model* (Müller-Crepon, Schvitz and Cederman 2023) to the case of the partitioning of voters into parties. The model was originally developed to partition geographical space into state territory. We transform the understanding of space to extend to a multi-dimensional electoral space. To apply the model we need to understand voters as the nodes of a network which is divided into partitions (candidates or parties) based on dyadic differences and similarities between voter characteristics, in short, electoral cleavages.

#### Probabilistic Partition Model

Following Müller-Crepon, Schvitz and Cederman (2023), we model the partitioning of voters as a Boltzman distribution

$$Pr(P = p_k) = \frac{e^{-\epsilon_k}}{\sum_{k=1}^{|\mathbb{P}|} e^{-\epsilon_k}}, \tag{A14}$$

where the chance that a given partitioning $p_k$ is realized decreases with its "energy" $\epsilon_k$. This energy can be interpreted as political tensions in a given division of voters into parties: the more voters are dissatisfied with the party they vote for in a given partitioning, the higher the tension and the less likely the partitioning emerges. Partitionings' energy $\epsilon_k$ results from attractive and repulsive forces $\epsilon_{i,j}$ between voters $i$ and $j$. These forces are only realized when $i$ and $j$ support the same party ($\mathbb{1}_{i,j} = 1$) and not otherwise:

$$\epsilon_k = \sum_{i,j \in L} \mathbb{1}_{i,j} \epsilon_{i,j}, \tag{A15}$$

$$\epsilon_{i,j} = \beta_0 + \gamma \, \mathbf{x}_{i,j}, \tag{A16}$$

The political attraction and/or repulsion between pairs $L$ of voters $i$ and $j$ is determined on the one hand by a constant baseline attraction $\beta_0$, as well as a vector of dyadic comparisons $\mathbf{x}_{i,j}$ between them. These comparisons can include binary indicators of differing ethnicity or gender, as well as distance measures, such as their wealth difference or geographic distance between them. Intuitively, we expect individuals with different ethnic backgrounds or vastly different incomes to be less likely to vote for the same party – indeed, were the same party trying to attract them, it might end up not succeeding or splitting. The vector of $\gamma$ parameters indicates the effect of each dyadic voter comparison on the attraction and repulsion between voters and thus ultimately the partitioning of voters into parties. Estimating parameters in $\gamma$ is therefore our ultimate goal.

As can be seen, similar to our setup in the baseline analysis, this formulation of vote choice is entirely dependent on comparisons between voters and does therefore not pre-suppose the existence of any party or set of parties. These emerge endogenously as the result of co-voting between voters. This allows for estimating the model across countries or country-periods with differing sets of parties and candidates.

We estimate Eq. A16 using the same data as used in the main analysis. In fact, the set of dyadic comparisons constructed for each country-round of the Afrobarometer can be represented as a graph $G_{c,t}$ of voters $i, j \in N$ who are associated with party or candidate preferences. The edges $L$ of $G$ encode the covariates $\mathbf{x}_{i,j}$ in Eq. A16 that determine whether voters $i$ and $j$ are likely to vote for the same (attraction) or two different (repulsion) parties. These co-variates are the same as used in the main analysis. Instead of a separate fixed effect for each country, we add one variable which stores the average attraction between nodes from each country. This is derived as the intercept of an otherwise empty model estimated separately for each country.

**Results**

Table A23 presents the main estimates from the partition model, derived – as in the main analysis – from unconditional and conditional models of the effect of share mother tongues on respondents' joint support for presidential candidates and legislative parties. We find relatively large estimates which are stable across specifications and outcomes and associated with little uncertainty. Importantly and as in our main analysis, the effect associated with a shared mother tongue does not significantly change with the introduction of other covariates

Table A23: Shared mother tongue and respondents' partitioning into candidates and parties

| | Presidential candidates | | Party support | |
|---|---|---|---|---|
| | (1) | (2) | (3) | (4) |
| Constant | 0.0000 | 0.0044* | 0.0003 | 0.0036 |
| | [−0.0008; 0.0013] | [0.0015; 0.0069] | [−0.0003; 0.0025] | [−0.0003; 0.0062] |
| By-country intercept | 0.8142* | 0.8172* | 0.8466* | 0.8690* |
| | [0.7190; 0.9533] | [0.6911; 1.0118] | [0.7889; 0.9879] | [0.7819; 1.0359] |
| Shared mother tongue | −0.0091* | −0.0087* | −0.0095* | −0.0092* |
| | [−0.0119; −0.0063] | [−0.0116; −0.0057] | [−0.0162; −0.0057] | [−0.0155; −0.0059] |
| Shared religion | | −0.0004 | | −0.0007 |
| | | [−0.0027; 0.0008] | | [−0.0040; 0.0013] |
| Age similarity | | −0.0001 | | −0.0003 |
| | | [−0.0004; 0.0001] | | [−0.0007; 0.0001] |
| Shared gender | | −0.0066* | | −0.0043 |
| | | [−0.0089; −0.0013] | | [−0.0087; 0.0023] |
| Shared education | | −0.0022* | | −0.0018* |
| | | [−0.0035; −0.0011] | | [−0.0030; −0.0007] |
| Wealth similarity | | −0.0003* | | −0.0002 |
| | | [−0.0005; −0.0000] | | [−0.0005; 0.0001] |
| Shared occupation | | −0.0010 | | −0.0020* |
| | | [−0.0015; 0.0000] | | [−0.0028; −0.0012] |
| Geographic proximity | | −0.0012 | | −0.0011 |
| | | [−0.0038; 0.0000] | | [−0.0054; 0.0017] |
| Shared urban vs. rural | | −0.0018* | | −0.0015* |
| | | [−0.0028; −0.0010] | | [−0.0033; −0.0001] |
| Countries | 26 | 26 | 28 | 28 |
| Respondents | 16833 | 16833 | 12989 | 12989 |
| Edges | 5805059 | 5805059 | 3318087 | 3318087 |

*Notes:* 95% confidence intervals from country-level bootstrap in parenthesis. * Statistically significant at the 95% level.

We take two additional steps to gauge the comparability between the results from the partition model with our main results. First, we sequentially decrease the connectivity in graph $G$ that underlies the model to the point where each respondent is connected only to one other respondent. We ensure that each respondent is connected to a uniform number of edges by constructing the sparse graphs $G$ as the union of ring graphs. Each ring graph contains the full set of respondents in a random order and connects each respondent to their two ring-neighbours. Respondent orders are sampled such that the ring graphs do not contain overlapping edges.

Estimates from the disjoint graph of one dyad per respondent without any overarching network structure (see Figure A20) are very close to the estimates obtained from a logistic regression model using the main specification (see Table A21 above). The partition model indeed reduces to a simple edge-wise logistic regression where edges are not connected to each other. Deviations can be explained by the sampling error incurred when sparsening the graph.

Once the density of the network increases, coefficient estimates naturally decrease – this reflects that the importance (or 'energy') of any one edge in influencing the partition membership of each node decreases with the number of its edges. Yet, Figure A20 shows that the *ratio* between the effect of shared mother tongue and the remaining coefficients remains remarkably stable. We take this as further evidence that the network estimator closely mirrors our main results.

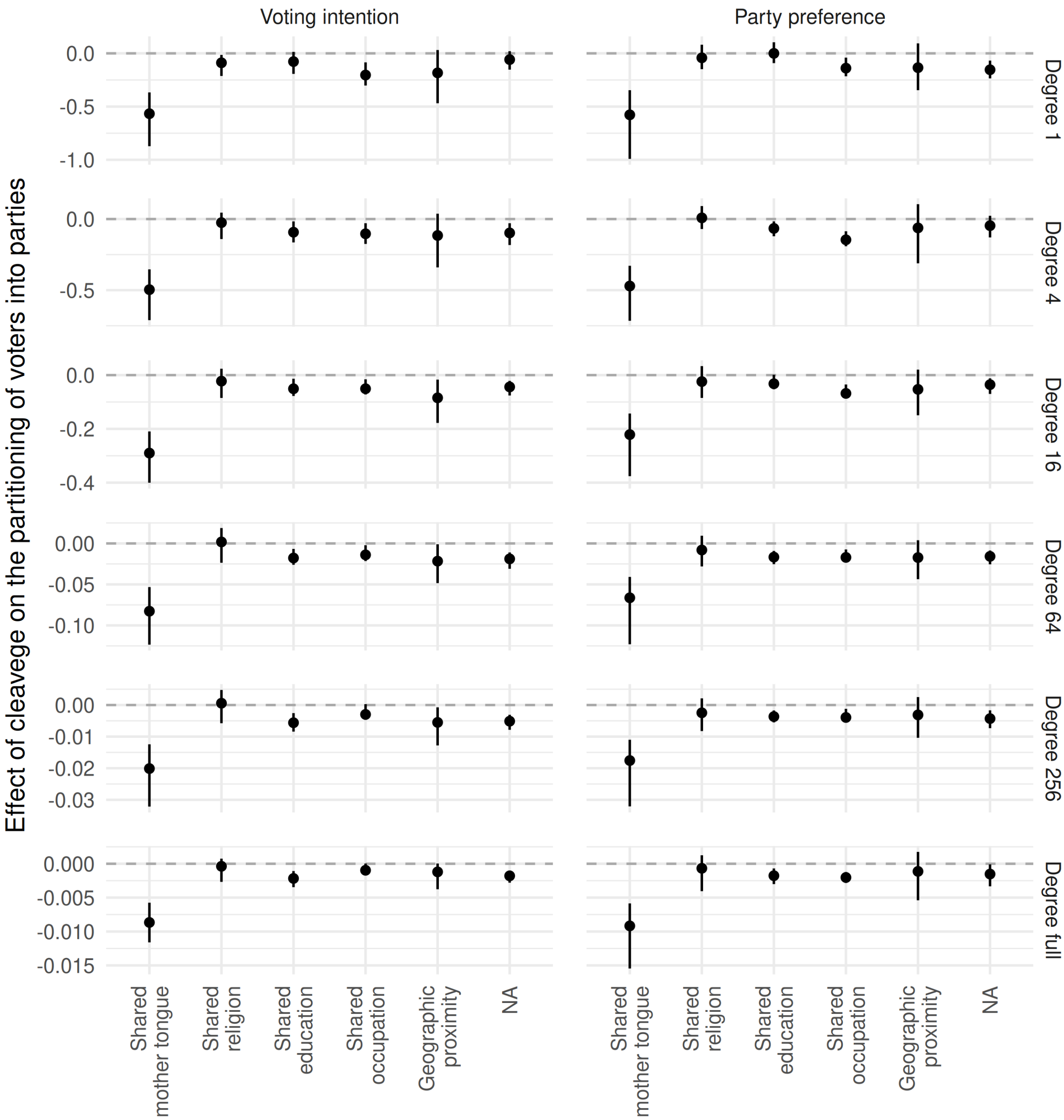



Figure A20: Partition model estimates by degree of connectivity

Note: The figure plots the most important predictors of the partitioning of respondents into presidential candidates and parties, by degree of network connectivity. Each set of coefficients results from estimating Eq. A16 using the full set of control variables. The estimated models are thus equivalent to Models 2 and 4 in Table A23 in all respects except for variation in network connectivity.

# G Country-level estimates

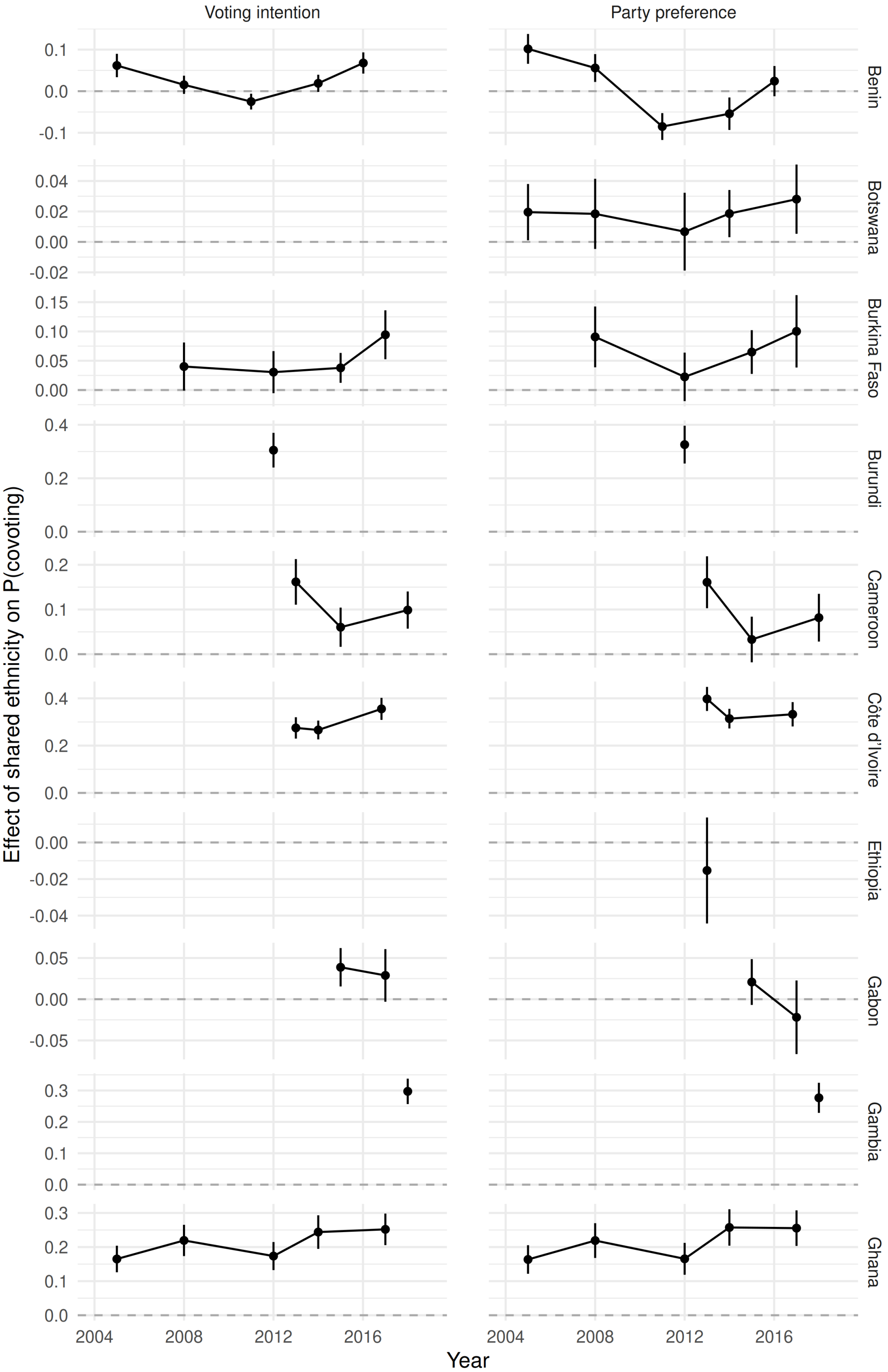



Figure A21: By country, over time, I

Note: Coefficients result from the fully specified model in Eq. 10 estimated separately for each Afrobarometer survey round since Round 3 with respondents' self-identified ethnicity as ethnicity indicator. The estimated models are thus equivalent to Models 2 and 4 in Table 1 in all respects except for the sample which varies by country-survey round.

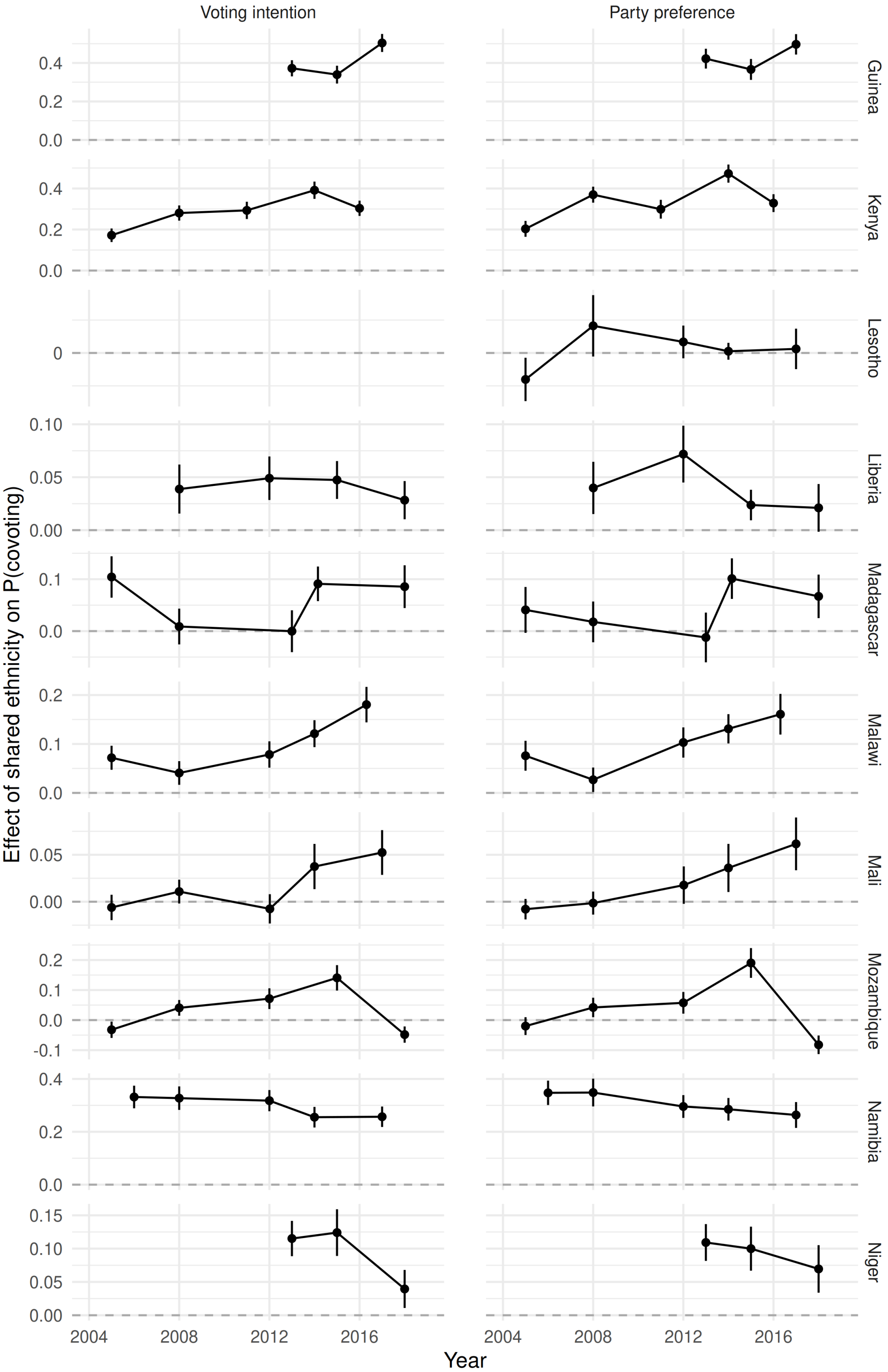



Figure A22: By country, over time, II

Note: Coefficients result from the fully specified model in Eq. 10 estimated separately for each Afrobarometer survey round since Round 3 with respondents' self-identified ethnicity as ethnicity indicator. The estimated models are thus equivalent to Models 2 and 4 in Table 1 in all respects except for the sample which varies by country-survey round.

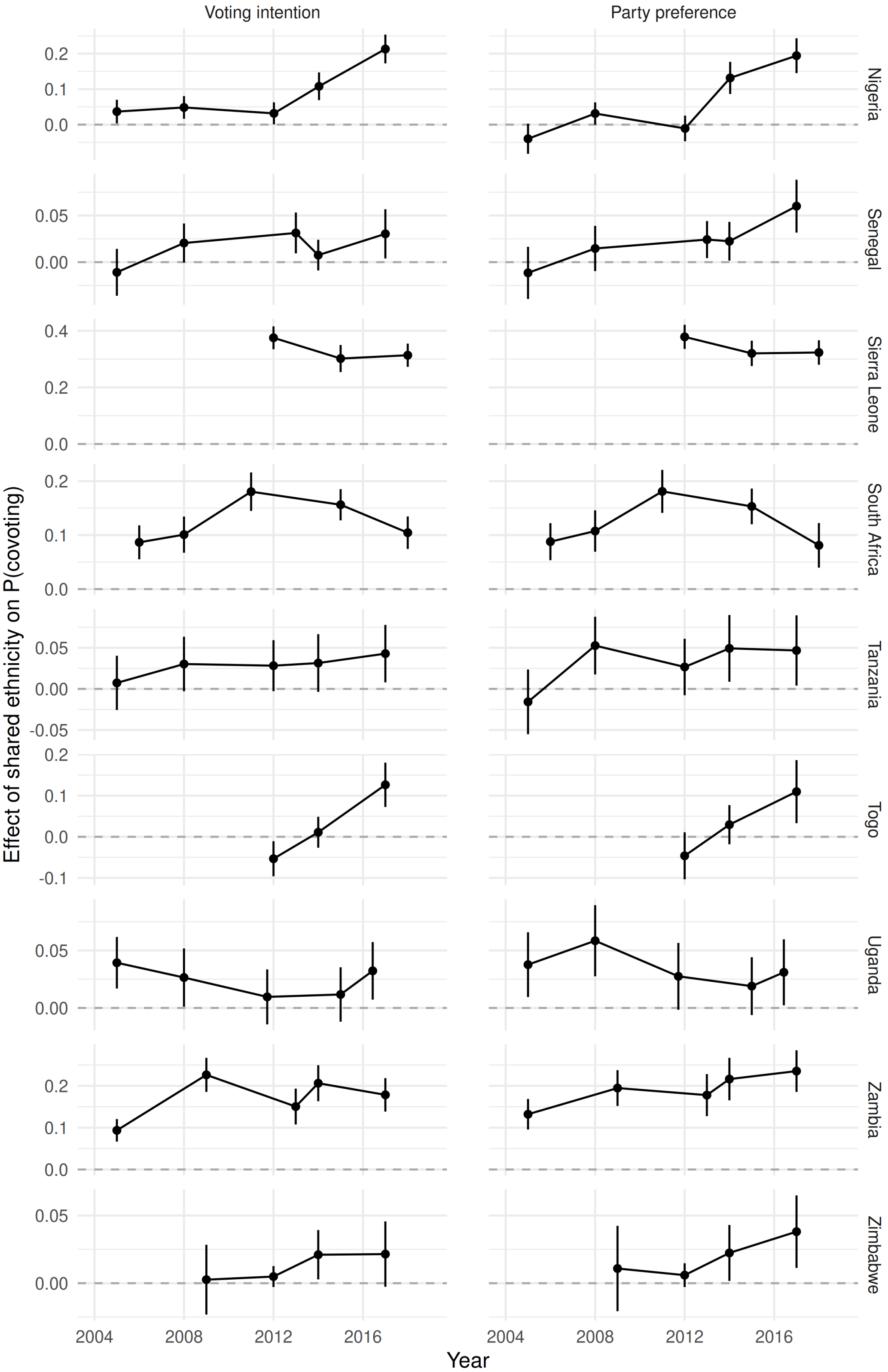



Figure A23: By country, over time, III

Note: Coefficients result from the fully specified model in Eq. 10 estimated separately for each Afrobarometer survey round since Round 3 with respondents' self-identified ethnicity as ethnicity indicator. The estimated models are thus equivalent to Models 2 and 4 in Table 1 in all respects except for the sample which varies by country-survey round.

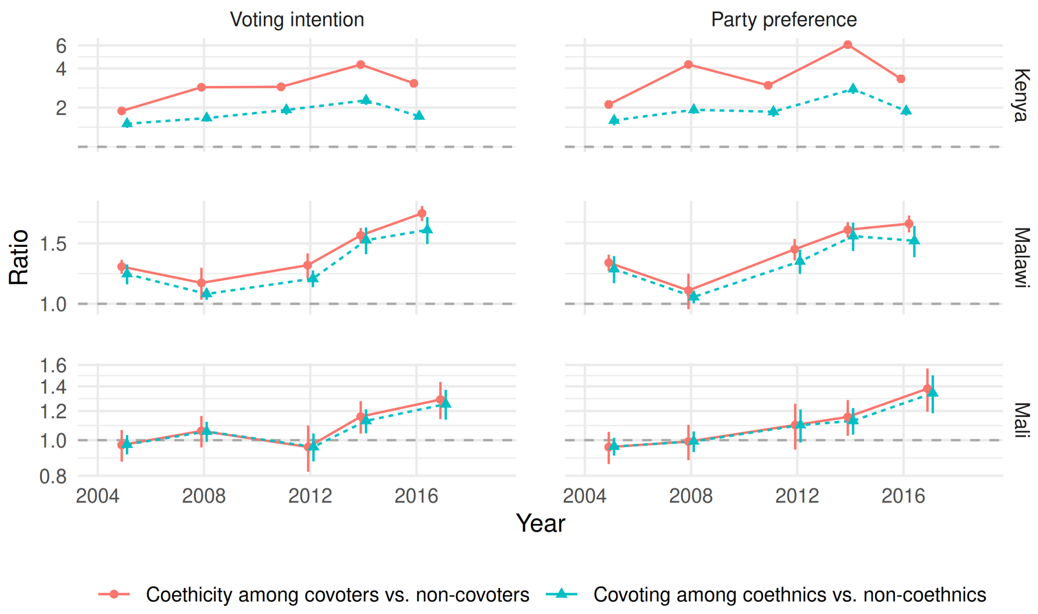



Figure A24: Group and Party driven changes in coethnic covoting in selected cases

Note: Odds ratios are computed by dividing the covariate-adjusted rate of coethnicity among covoters by the rate of coethnicity of non-covoters (red) and the rate of covoting among coethnics divided by the rate of covoting among non-coethnics (blue).

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