What matters to voters? Identifying important ideological dimensions with partial rankings of parties

Nina Nikiforova*

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Abstract

I propose a method to identify which ideological dimensions best rationalize voters' choices in a spatial voting framework using partial party rankings. Using the German Longitudinal Election Study, 2013–2022, and expert-based party placements, I show that the Left–Right dimension provides the best one-dimensional fit in every year, and that the economy–culture pair provides the best two-dimensional fit. Yet the fit of both models declines over time, while the anti-elite dimension gains explanatory power and by 2022 rivals the leading policy dimensions. These patterns suggest the growing importance of non-traditional ideological dimensions and other determinants of voting.

*Bocconi University. E-mail: nina.nikiforova@unibocconi.it. I am grateful to Guido Tabellini, Jérôme Adda, Francesco Decarolis, Catherine De Vries, Ro'ee Levy, Eduard Osipov, Cailin Slattery and Francesco Trebbi for their helpful comments and suggestions.

1 Introduction

What underlies voters' decisions at the ballot box? Answering this question is crucial for interpreting contemporary political landscapes, for informing policy responses, and for understanding party competition. For example, understanding the forces behind voters' recent realignment toward populist movements in Western democracies is essential for designing effective responses and projecting future developments.

In this paper, I show that partial rankings of parties can offer new insights into this question. I adopt a spatial voting framework in which voters choose the party closest to their ideal point in a policy space. Positioning parties along alternative dimensions (economic, cultural, environmental), or pairs of dimensions, makes it possible to determine whether reported partial rankings are consistent with spatial voting along those dimensions. In other words, this approach uses partial rankings to identify the ideological lines of voters.

To conceptualize the problem, I apply revealed-preference logic to stated rankings and adapt the setup of Degan and Merlo (2009). The authors show that the hypothesis that a voter votes ideologically is falsifiable only with individual-level data from several simultaneous elections. They demonstrate that any voting profile (the set of an individual's voting choices) is rationalizable in the absence of such data, whereas with data on simultaneous elections a given profile can be shown to be non-rationalizable. I treat partial party-ranking questions as a sequence of simultaneous elections: at each step, after removing previously selected parties, the next position is assigned to the party that would "win" among the remaining alternatives. The partial ranking itself corresponds to the voting profile.

I apply this framework to the German Longitudinal Election Study (GLES) from 2013–2022, which records each respondent's partial ranking of the three most-preferred parties (*triples*).[1] I complement these data with expert-survey estimates from Jolly et al. (2022) on party positions along several dimensions – Left-Right, economy[2], anti-elite (anti-establishment)[3], culture, EU (European Union), environment, and religion. The main exercise consists of positioning parties separately on each of these one-dimensional scales and computing, year by year, the share of reported triples that are rationalizable along each dimension. This exercise relies on a spatial voting framework in which voters choose the party closest to their ideal point.

[1]From 2013 onward, I can construct identical choice sets across years because Alternative for Germany (AfD) first appears among the answer options in 2013; 2022 is the last year available.

[2]Economy denotes the *Left-Right economic* scale, while Left-Right denotes the *Left-Right Ideology* summary scale. The economic scale is narrower, capturing only parties' economic policy positions; Left-Right also captures social and cultural elements of proposed policies.

[3]Substantively, the anti-elite dimension captures opposition to established elites broadly conceived – mainstream parties, technocrats and experts, media, and supranational institutions, and extends to support for relaxing the checks and balances that constrain executive power in the name of popular sovereignty (Mudde, 2004; Guriev and Papaioannou, 2022). The CHES item measures the salience of anti-elite rhetoric in a party's public stance, which I treat as the strength of these positions. One might object that this dimension is not tied to a clearly defined set of policies and should therefore not be considered a policy dimension. Views differ: Bellodi et al. (2025) suggest that establishment-sceptical parties advance distinct policy bundles, while classic definitions of populism treat anti-elitism primarily as rhetoric (Mudde, 2004). I refer to the anti-elite dimension as a policy dimension reflecting the conflict over elites and institutional constraints, as just defined.

There are two main findings from the one-dimensional setup. First, the Left-Right dimension has the greatest explanatory power: it rationalizes the largest share of observed partial rankings (*triples*) in every year, though its share has declined from above 50% in 2013 to 40% in 2022. Moreover, I show that in most years the Left-Right dimension provides the best possible fit — in the sense that no alternative dimension could rationalize a larger share of observed triples. This finding suggests that in most years *the Left-Right dimension is the best attainable one-dimensional model of voters' behavior*.

This result is important for two reasons: it points to the paradigmatic Left-Right dimension as the most important not only among the dimensions considered, but, in most years, among all possible one-dimensional alternatives; and it quantifies the explanatory power of a dimension that is widely used but whose empirical scope has remained unclear. The fit for the economy dimension is slightly worse than for the Left-Right dimension throughout the study period; the one-dimensional fit for the cultural dimension is weak.

Second, the anti-elite dimension's explanatory power rises sharply and, by 2022, it rationalizes the same share of observed partial rankings as the economy dimension. This suggests that *voters increasingly perceive the anti-elite conflict as a dimension of party competition*. I show that these results are not confined to the voters of distinctively anti-elite parties, but reflect a broader phenomenon across the whole electorate. This resonates with recent developments in Germany: as the anti-elite far-right Alternative for Germany (AfD) has grown, about three quarters of Germans view the AfD as a threat to democracy and are concerned about the consequences its rise could bring (Forschungsgruppe Wahlen, 2024; Infratest dimap, 2025). That anti-elite attitudes structure the choices of populist supporters is unsurprising; that they also structure the choices of mainstream supporters is more striking and suggests that pro-elite voters, in the presence of a strong enough populist threat, might decline to trade away their pro-elite commitments at the ballot box.

I also show that these results are not mechanical, as the rationalizability of partial rankings aligns with voters' stated views. Respondents whose triples are rationalizable along a given dimension are also more likely to name a "most important problem" aligned with that dimension. For example, those whose triples are rationalizable along the economy dimension more often cite the economic situation, social benefits, unemployment, or taxes as the most important problem.

I then ask whether one can formally assess which dimensions of party competition matter most to voters. I formulate testable hypotheses: a dimension "matters" if, under ideological voting, it explains the observed partial rankings across the population better than a random benchmark. I construct this benchmark by randomly drawing party positions 100,000 times and computing, for each draw, the share of rationalizable triples. The null of no ideological voting along a given dimension is rejected at conventional significance levels ($p \leq 0.05$) for the Left–Right and economic dimensions in all years, and for the anti-elite dimension in recent years. These results confirm the statistical validity of the findings discussed above.

I show that the test of ideological voting in one dimension is robust to misreporting of party preferences due to stigmatization. Because the AfD (the most stigmatized party) is at the extreme of all considered ideological dimensions, a respondent with well-formed policy views who omits AfD but otherwise ranks parties ideologically will not change the test's conclusion. By contrast, such misreporting would be problematic for recovering voters' bliss points from partial rankings.

Finally, I extend the analysis to two dimensions by asking which pair of dimensions, and with what relative weights, rationalizes the largest share of observed triples in each year. The results identify a stable best-fitting pair, economy and culture, with economy receiving the larger weight (60%) in all years except 2022. This closely mirrors the one-dimensional findings, since the Left–Right scale is often interpreted as combining economic and cultural conflict. At the same time, the fit of the best specification declines over time, while the Left–Right/anti-elite pair improves relative to other pairs and ranks among the top specifications by 2022. Together, these patterns suggest the growing importance of non-traditional dimensions and other factors in voting choice.

The paper proceeds as follows. Section 2 overviews related literature. Section 3 presents the setup adopted from Degan and Merlo (2009). Section 4 formalizes the test for identifying important dimensions of party competition. Section 5 describes the data. Section 6 illustrates how partial rankings help identify ideological lines. Section 7 reports the main results and discusses different aspects of the analysis. Section 8 concludes.

2 Related Literature

The contribution of this paper is twofold: a novel method to recover the ideological lines of voters, and new evidence on the evolution of the German political landscape over 2013–2022, which speaks to the literature on political competition in Western democracies, in particular in the presence of populism.

Methodology: recovering the ideological lines of voters. Recovering the ideological lines of voters requires understanding individual preferences and a model that maps preferences into voting choices. While this paper proposes using partial party rankings and a spatial voting model, the literature has approached this task in at least three other ways.

First, there is a body of literature that relates voting choices to stated preferences (Fiorina and Abrams, 2008), with only a few papers embedding stated preferences in an explicit mapping from preferences to choices (Ansolabehere and Puy, 2018; Danieli et al., 2023). A related literature recovers latent ideologies or groups of like-minded voters from stated preferences across a broad range of topics (Uscinski et al., 2021; Draca and Schwarz, 2024; Desmet et al., 2025). Stated preferences, however, are not necessarily truthful and might encode partisan affiliation (Graham, 2026), which is especially problematic when trying to recover the ideological lines of voters.[4]

Second, another strand connects ideological lines and voting choices to answers to *the most important problem facing society* question (Green and Hobolt, 2008). As Wlezien (2005) argues, this approach conflates the salience of a problem with its scale; see Kuziemko et al. (2015) for an example, and Sides et al. (2023) for evidence that these answers need not predict vote choice.

Third, a growing literature estimates structural models of voting, fixing the ideological dimensions ex ante (Henry and Mourifié, 2013; Merlo and De Paula, 2017; Iaryczower et al., 2024; Longuet-Marx, 2025). While this literature aims to recover the full distribution of preferences over predefined dimensions, this paper focuses solely on recovering the ideological lines, leaving the estimation of the full distribution of preferences over the recovered dimensions for future work.

Relative to these strands, this paper (1) explicitly builds on the spatial voting model (Downs, 1957), (2) focuses solely on recovering ideological lines, and (3) uses partial party rankings rather than other sources of data. Methodologically, it is closest to Degan and Merlo (2009), who establish the falsifiability of ideological voting with data on simultaneous elections but do not offer a framework to recover the ideological lines. Treating partial party rankings as simultaneous elections is what makes the approach feasible: individual-level data on simultaneous elections are rarely available, whereas partial rankings are elicited in many surveys.

The paper is also related to Apesteguia and Ballester (2025), who study the rationalizability of survey responses. While they search for the best-fitting one-dimensional order of questions, the main analysis in this paper takes candidate orders from existing policy dimensions and computes the shares of individuals whose responses are consistent with each.

Results: recovered ideological lines of voters. The paper contributes to the literature on the recent evolution of political cleavages (De Vries et al., 2013; Noury and Roland, 2020; Bonomi et al., 2021; Gethin et al., 2022; Margalit et al., 2025; Danieli et al., 2023; Gennaioli and Tabellini, 2025). The results justify the literature's focus on the Left-Right dimension and attach a number to it: the share of the electorate whose preferences are consistent with this dimension is about 40–50% in my sample. They also speak to the debate over the relative importance of economic and cultural cleavages in Western democracies, showing that the recovered ideological lines are primarily economic, with cultural conflict entering as an important component only when paired with the economic dimension.

The finding regarding the anti-elite dimension deserves special attention. Anti-elite sentiment is a

[4] Whenever an individual reports a position on an issue close to the party they vote for, this could be for two opposite reasons: (1) the issue matters most to the individual, and they have chosen the party because of its position on it; or (2) the issue is unimportant, and the individual has shifted their stated position towards the party's out of conformism.

core feature of populism, and a broad literature documents the anti-elite preferences of populist parties' voters (Algan et al., 2017; Rooduijn, 2018; Guriev and Papaioannou, 2022); by contrast, the literature on the ideological lines of voters in the presence of populism is scarcer (Noury and Roland, 2020). This paper shows that the anti-elite dimension has become an increasingly important ideological line in society as a whole—not merely that voters of anti-elite parties hold anti-elite preferences or attach weight to this dimension. The closest findings are those of Uscinski et al. (2021) and Draca and Schwarz (2024), who recover the ideological structure of the electorate from voters' attitudes and find that voters are best organized along a Left-Right dimension and a dimension characterizing their anti-elite attitudes. This paper is different in that it focuses solely on recovering separate ideological lines, and does so in a way more directly related to vote choices: from party rankings rather than issue attitudes.

3 Baseline Model

The formalization in this section borrows from Degan and Merlo (2009). Consider a situation in which N voters are facing m simultaneous elections. Assume an ideological space $Y = \mathbb{R}^k, k \geq 1$. For the election $e \in \{1, \dots, m\}$, let J_e denote the candidate set and $q_e = |J_e|$ the number of candidates in election e. Each candidate j is characterized by a unique position in the ideological space $y^j \in Y$, known to voters.

Each voter $i \in \{1, \dots, N\}$ has an ideal policy position $y^i \in Y$. Under a spatial voting model, voter i's preferences over candidates in election e are given by $U_e^i(j) = u(\text{dist}(y^i, y^j))$, $j \in J_e$, where u is a decreasing function, and $\text{dist}(\cdot, \cdot)$ denotes a distance function. Thus, i prefers candidate j_e to l_e in election e if $\text{dist}(y^i, y^{j_e}) < \text{dist}(y^i, y^{l_e})$.[5]

For each voter i let $v^i = (v_1^i, \dots, v_m^i) \in V^m$ denote her voting profile, where $v_e^i \in J_e$ denotes her voting choice in e, and $V^m = J_1 \times \cdots \times J_m$ is the set of all possible distinct voting profiles in m elections. Hence, i's voting profile contains the list of candidates she votes for in the m elections.

Next, Degan and Merlo (2009) introduce the concepts of *ideological voting* and *a voting profile rationalizable by ideological voting.*

Definition 1. Voter i votes ideologically in election e if she votes for a candidate whose position is closest to her bliss point: $\text{dist}(y^i, y^{j_e}) < \text{dist}(y^i, y^{l_e})$ for all $l_e \in J_e$ with $l_e \neq j_e$ implies that $v_e^i = j_e$. Voter i *votes ideologically* if she votes ideologically in all elections $e = 1, ..., m$.

Definition 2. A voting profile $v \in V^m$ is *rationalizable by ideological voting* if there exists some non-empty subset of the ideological space, $Y^v \subseteq Y$, such that if $y^i \in Y^v$ and i votes ideologically, then $v^i = v$.

[5] Note that this formalization assumes a symmetric utility function around the bliss point, which is standard in the political economy literature (Persson and Tabellini, 2002). I relax this assumption for some of the results below.

In practice, rationalizability is verified by checking a finite system of inequalities. See Section C for more details. Building on this setup, I now turn to the test of ideological lines of voters.

4 Test of Important Dimensions of Voting

Building on the setup introduced in the previous section, this section develops a statistical approach to test for important dimensions of party competition.

Let $Y_d = \mathbb{R}^{|d|}$ denote the policy space spanned by dimensions $d = \{d_1, d_2, \ldots\}$, where $|d|$ is the number of dimensions considered. The canonical example of a dimension in political economy is the Left–Right dimension, and both voters and parties are assumed to have positions along this dimension.

Given Definition 2, one can measure the share of voters whose profiles are consistent with ideological voting along d. Say that a profile $v \in V^m$ is *rationalizable along* d if it is rationalizable by ideological voting on the ideological space Y_d, where voting ideologically along d means that distances are computed between positions on d (or dimensions, if $|d| > 1$), y_d and y_d^j. Denote by S_d the set of voters whose observed profiles are rationalizable along d:

$$S_d = \big\{\, i \in \{1, \ldots, N\} : v^i \text{ is rationalizable along } d \,\big\},$$

and define the *share of profiles rationalizable along* d as

$$s_d = \frac{|S_d|}{N}.$$

This is the key metric used throughout the paper. A natural question is whether differences in s_d are meaningful or could arise by chance. I now formulate a test to distinguish between the two.

I formulate the null and alternative hypotheses for a given Y_d as follows:

H_0: *Voters vote ideologically, but not along dimension* d,

H_1: *Voters vote ideologically along dimension* d.

A key maintained assumption is that voters choose triples ideologically – that is, consistent with a spatial voting model along some dimension. The test then identifies the dimensions along which their choices are most consistent.

The test statistic is the share of rationalizable voting profiles, s_d, and the decision to reject H_0 is based on the probability, under H_0, of observing a result at least as extreme as s_d. To calculate this probability, I first randomly draw party positions d' ($|d'| = |d|$) from a uniform distribution[6] over the

[6] An important assumption here is that parties' positions are iid. I partly mitigate this concern by using conservative significance cutoffs in subsequent steps. To address the concern that the uniform distribution might yield an overly permissive test, I also provide a robustness exercise in which I repeat the analysis with reshuffled party positions.

bounded policy space $Y_{d'}$[7]. Next, I calculate the share of rationalizable voting profiles $s_{d'}$, assuming ideological voting. Repeating this exercise yields a distribution of $s_{d'}$. I reject H_0 at level α if $\mathbb{P}(s_{d'} \geq s_d) \leq \alpha$.

Intuitively, rejecting H_0 implies that the share of voting profiles rationalized by Y_d is much higher than the share rationalized by a random $Y_{d'}$. Note that, given the definition of rationalizability by ideological voting, there is no direct mapping from Y_d to s_d. Hence, the simulation-based hypothesis test described in this section provides a practical approach.

Partial Party Ranking. To implement the test empirically, two ingredients are required: (i) individual voting choices in simultaneous elections and (ii) candidate positions. In practice, such individual-level data are difficult to obtain, since simultaneous elections with distinct candidate sets are rare and voting data are sensitive. I therefore use survey-based partial party rankings: respondents repeatedly select their preferred party from a set, with the chosen party removed after each choice. Each selection can be interpreted as an election with a different candidate set. I obtain candidate positions from expert surveys.

5 Data

The analysis covers the period from 2013 to 2022 and focuses on a single country, Germany. Germany is a multiparty democracy with, historically, two main establishment parties: the centre-left SPD (Social Democratic Party) and the centre-right CDU (Christian Democratic Union), and several smaller parties: the Left, the Greens, and the FDP. The most important recent development in the German political landscape is the emergence and rise of the radical-right anti-establishment party Alternative for Germany (AfD). The party narrowly missed the 5% electoral threshold for entering the Bundestag in 2013, and entered the Bundestag in the 2017, 2021, and 2025 elections with 12.6%, 10.4%, and 20.8% of the vote, respectively. The analysis in this paper uses data on partial rankings and positions of six parties: AfD, CDU, FDP, Greens, the Left, and SPD.

In this section, I describe the two data sources: (i) the German Longitudinal Election Study (GLES), which elicits partial party rankings, and (ii) an expert survey on party positions across different dimensions.

5.1 Voter survey

The first dataset is the German Longitudinal Election Study (Forschungsgruppe Wahlen, 2023). The dataset consists of annual surveys of a nationally representative sample of the German electorate that focus on the opinions and attitudes of the voting population in the Federal Republic on current political topics, parties, politicians, and voting behavior. Each wave includes roughly 20,000–30,000 respondents.

[7] In practice, I restrict the policy space to the interval $[0, 10]$, matching the scale of the data on party positions.

The data include demographic characteristics of respondents, as well as answers to some political questions. My main focus is on the questions: *"Which of these parties do you like best?"*, *"And which one second best?"* (asked twice). These questions are suitable for the purpose of this paper for at least two reasons. First, these questions could motivate the sincere party choice model. This is in contrast to questions like "If the election were today, which party would you vote for?", which tend to elicit more strategic responses.[8]

Second, respondents are given a clear choice set to choose from, which mitigates concerns regarding the unobservability of the consideration set. Figure 1 shows the questions that a hypothetical respondent who answers *1. SPD, 2. Greens, 3. Left* would see. This setup allows a clear mapping of partial ranking to simultaneous elections. The final dataset contains a total of about 170,000 respondents for the years 2013–2022.[9]

Which of these parties do you like best?	**And which one second best?**	**And which one third best?**
☐ CDU / CSU	☐ CDU / CSU	☐ CDU / CSU
☒ SPD	☐ AfD	☐ AfD
☐ AfD	☒ Greens	☐ FDP
☐ Greens	☐ FDP	☒ Left
☐ FDP	☐ Left	☐ NA
☐ Left	☐ NA	
☐ NA		

Figure 1: Example of choice situations

5.2 Party Positions

The second dataset is the Chapel Hill Expert Survey (CHES) (Jolly et al., 2022), which measures party positions across multiple ideological dimensions. Guided by the literature, I focus on Left–Right, economy, culture, anti-elite, EU, environment, and religion (see, e.g., Wheatley and Mendez, 2021).

Table A1 reports the corresponding CHES scores. CHES places parties on each dimension on a 0–10 scale. I now discuss in more detail the dimensions that prove most relevant for this paper. On the economy dimension, higher values indicate stronger support for free-market policies such as lower taxes and deregulation, whereas lower values reflect interventionist positions favoring redistribution and a larger role for the state. For example, AfD (8.3) and FDP (8.0) are highly market-oriented, while the Left (1.3) favors extensive state intervention. On the culture dimension, higher values correspond to greater

[8] Respondents remain unlikely to choose parties polling far below the 5% hurdle. For this reason the NPD is dropped for all the years in which it was present in the choice set. This affects about 1% of all observed triples.

[9] All observations with fewer than three party-choice answers are dropped, as three answers are needed to falsify ideological voting in two dimensions. Fewer than three answers occur when individuals either refused to respond or indicated that they would abstain. In addition, the CSU is dropped from the choice set and all triples containing it are excluded. The survey presented respondents with an unrealistic choice between the CDU and the CSU — two parties that form a single federal-level alliance (the Union) and never compete in federal elections — which could distort the responses of those who perceive the CSU as a genuine choice option (in particular, Bavarian voters, for whom the CSU is the CDU's counterpart); indeed, some observed triples include both the CDU and the CSU, which is infeasible.

cultural conservatism (e.g., AfD at 8.7), while lower values indicate cultural liberalism (e.g., Greens at 2.2). The Left–Right Ideology dimension is a broader summary measure that is often interpreted as combining parties' economic and cultural stances. The anti-elite dimension captures the extent to which parties depict political elites as corrupt, untrustworthy, or detached from ordinary citizens: AfD scores 7.8, consistent with a strong anti-elite framing, whereas mainstream parties such as CDU (0.8) and FDP (0.9) exhibit little anti-elitism.

6 Illustration: partial party rankings

This section illustrates the usefulness of partial party rankings for recovering the ideological lines of voters. I first present the substitution patterns from first to second choices and discuss the immediate conclusions they support. I then illustrate the recovery of ideological lines using the concept of triple rationalizability along a dimension, the main object of interest in the remainder of the paper.

6.1 Second-choice data

The use of second-choice data, while it has proven useful in the applied industrial organization literature (Berry et al., 2004; Conlon and Mortimer, 2021), has not yet found broad application for preference estimation in the political economy literature. I now present the raw data on first- and second-best party choices, and discuss the immediate conclusions one could draw from them.

Figure 2 plots the joint distribution of first and second party choices – the share of respondents naming the row party as their first choice and the column party as their second choice (see also Figure A1, which normalizes these shares by first choices, yielding diversion ratios). The parties are ordered along the Left-Right dimension for ease of interpretation (the AfD is the furthest right, the FDP next, ..., the Left is the furthest left). Large values near the diagonal therefore imply that individuals substitute towards the parties closest to their preferred party on the Left-Right dimension. For instance, among all respondents, 6% choose the CDU as their first choice and the FDP as their second, which is consistent with those voters making their party choices along the Left-Right dimension. Likewise, the Greens–SPD pairing is consistent with voting along the Left-Right dimension. Overall, the data suggest that a large share of second choices cluster near the diagonal and can be explained by the Left-Right dimension.

But this is not true for many other pairs. The second-most common choice among AfD voters – the Left – is, in turn, not consistent with voting along the Left-Right dimension, but could be rationalized by another dimension. One such dimension is the anti-elite dimension. The analysis that follows employs a similar logic, but over the three best choices. I now turn to the illustration of this approach.

Figure 2: Distribution of first- and second-choice party pairs.

Notes: Each cell reports the share of respondents naming the row party as their first choice and the column party as their second choice. Shares are aggregated across the ten survey years (GLES 2013–2022); the 30 off-diagonal cells sum to 100%. Parties are ordered along the Left-Right dimension.

6.2 Triple rationalizability

In the analysis that follows I do not restrict attention to pairs (the two first-ranked parties), but focus on triples (the three first-ranked parties), as triples provide richer variation for determining the ideological lines of voters. The main analysis in this paper hinges on the concept of triples' rationalizability – i.e. whether a specific triple is rationalizable or not with respect to a given dimension.

Figure 3 illustrates the approach with three examples on the Left-Right dimension, under the assumption of a utility function symmetric around the bliss point—that is, at each step the voter chooses the party closest to her bliss point from the remaining choice set. This assumption serves as a benchmark for the main model used throughout the paper.

The triple FDP-CDU-AfD is rationalizable, as there is a set of bliss points that gives rise to this triple; likewise, SPD-Greens-CDU is rationalizable. The triple SPD-Greens-Left, by contrast, is not: from any bliss point that ranks SPD above the Greens, CDU is closer than the Left, so the Left cannot rank third. Note that the second and third triples share the first two choices and differ only in the third — yet one is rationalizable and the other is not. Rationalizability thus depends on the entire reported ranking, not only on the top choices: each additional choice imposes further restrictions on the set of consistent bliss points. Eliciting three choices rather than two therefore sharpens the identification of ideological lines, and deeper partial rankings would sharpen it further. I now turn to the main results: the rationalizability of observed triples along different dimensions and across years.

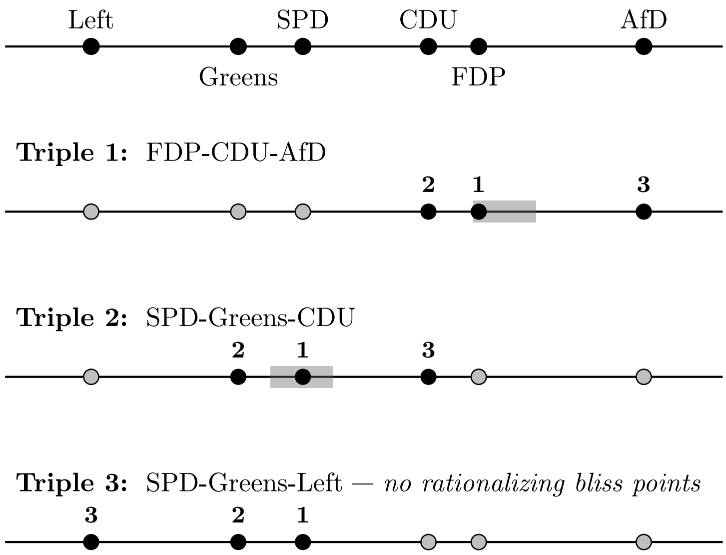

Figure 3: Bliss-point sets rationalizing selected triples on the Left–Right dimension

Notes: The top line shows the six parties ordered as on the Left–Right dimension (positions stylized for readability; exact CHES 2014 scores in Table A1). Each line below marks a reported triple, with numbers indicating the first, second, and third choice; the gray area is the set of bliss points y for which that triple arises under ideological voting with symmetric utility, i.e., for which the three chosen parties are, in order, the closest, second-closest, and third-closest of all six. For Triple 3 no such bliss point exists: ranking SPD above the Greens places y to the right of the Greens–SPD midpoint, but from any such point CDU is closer than the Left, so the Left cannot rank third.

7 Results

This section presents the main results of the paper. I first report the shares of rationalizable triples across dimensions and over time, and discuss the test of the important dimensions of party competition. I then turn to several aspects of the approach, including robustness to misreporting and the relation to stated most important problems.

7.1 Ideological lines of voters

In the survey data, individuals are asked about their three most-preferred parties ($m = 3$). In the following, I refer to voting profiles v^i of chosen parties as *triples*. My main object of interest is the share of observed *rationalizable triples* along different dimensions s_d and the inference about important dimensions of party competition that one can draw from it. I focus on the one-dimensional case ($|d| = 1$) first and discuss the two-dimensional case later.

Figure 4 presents estimates of the shares of rationalizable triples s_d across different dimensions d and years for one-dimensional models. For three of the seven dimensions considered (*culture*, *religion*, *environment*) the share of individuals whose triples are rationalizable by ideological voting along these dimensions is below 20% in every year. In contrast to these dimensions, about 40–50% of individuals choose triples consistent with ideological voting along the *Left–Right* dimension.

Hence, this exercise quantifies the possible scope of these dimensions: a broad literature in political economy uses the Left-Right and other dimensions to model the political behavior of voters, while the extent to which these dimensions are applicable remains unclear.

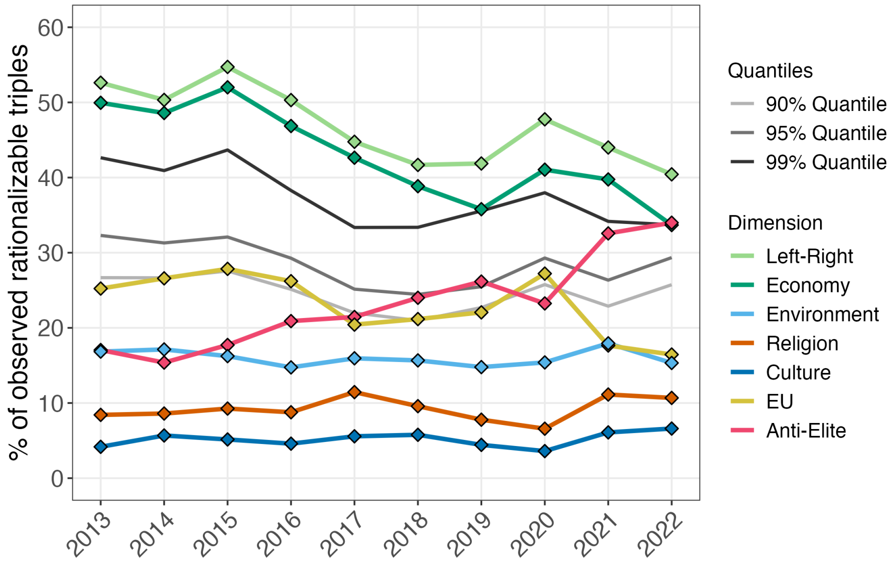

Figure 4: Share of rationalizable observed triples across years and dimensions.

Notes: The plotted values are the shares of rationalizable triples for one-dimensional models for different years and dimensions. Party positions fixed at CHES 2014 for all years. Gray lines denote 99%, 95%, 90% quantiles of the share of rationalizable triples based on 100,000 simulations of random party positions. *Note: Data collection in 2020 differed due to the COVID-19 pandemic (Glasenapp et al., 2024).*

There are clear time trends in the shares of rationalizable triples (fits) across dimensions. The fit declines over time for both the Left–Right and the economy dimensions. By contrast, the anti-elite dimension steadily gains explanatory power, reaching the economy dimension's fit by 2022. These results point to the central, but declining, role of the Left–Right dimension over time and the growing significance of the anti-elite dimension for voters.[10] Figure A2 formalizes the evolution of the two key dimensions by estimating a linear time trend in the shares of rationalizable triples.[11] The share rationalizable along the Left-Right dimension falls by 1.32 percentage points per year, while the share rationalizable along the anti-elite dimension rises by 1.94 percentage points per year; both trends are statistically significant at the 1% level.

The rise of the anti-elite dimension coincides with the rise of an anti-elite party, the AfD. The results are not, however, a mechanical artifact of that rise: they hold when the AfD is excluded from the analysis (Figure A3). Moreover, two clearly anti-establishment parties sit at the anti-elite end of this dimension, the AfD and the Left. Splitting respondents by their first-choice party shows that supporters of these two parties choose triples rationalizable along the anti-elite dimension at a consistently high rate throughout the period (≈30%, with a trend indistinguishable from zero), whereas supporters of the

[10] To assess sampling uncertainty, I bootstrap the yearly shares by resampling respondents with replacement within each year (keeping the observed yearly sample sizes, 1,000 replications) and recomputing the share of rationalizable triples per dimension; Figure A5 shows the resulting 95% percentile confidence bands. The intervals are narrow, and all results that follow are unaffected by sampling uncertainty.

[11] The trend is estimated at the individual level: a linear probability model regresses the indicator for whether a respondent's observed triple is rationalizable along the given dimension on a linear time trend (year − 2013), pooling all survey years 2013–2022. The fitted lines in Figure A2 weight each survey year by its number of respondents and therefore coincide with these individual-level estimates.

mainstream parties do so *increasingly*, starting from a much lower level (Figure A4). Since respondents do not choose the AfD or the Left consistently more often over time (see Table A2 for the election years), the aggregate rise of the fit of the anti-elite dimension cannot be attributed to a growing share of anti-elite party voters, and instead reflects mainstream-party voters whose rankings increasingly align with the anti-elite dimension.

This highlights that the anti-elite dimension has become a dimension of competition even among voters of pro-elite parties, and suggests that the nature of political conflict is shifting from conflict over the Left-Right dimension to conflict over the role of elites and democratic institutions. This could be read as a positive development for institutions and democratic stability: while the anti-elite ideological line is known to matter for supporters of populist parties, the fact that it also matters for supporters of mainstream parties suggests that, in the presence of a real threat, voters with pro-elite positions are unwilling to compromise on those values at the ballot box.

Note that the shares of rationalizable triples do not sum to 100% across dimensions, as a single triple can be rationalizable along several dimensions. The share of triples rationalizable along at least one of two dimensions therefore depends both on how much each dimension rationalizes separately and on the overlap between them. For instance, the share of triples rationalizable along the Left-Right or the anti-elite dimension reaches its lowest level, 47%, in 2017 and its highest, 58%, in 2022.[12]

One concern is that voters may not perceive party positions as experts do. Figure A6 suggests this concern is likely unfounded, as respondents' perceived Left-Right positions (the only dimension available in the survey) are highly correlated with expert ratings. Moreover, the results remain quantitatively similar when allowing for noise in party positions (Figure A11).

The results in this section rely on a utility function that is symmetric around the bliss point. Figure A7 reports the corresponding results when this symmetry assumption is relaxed. Under a non-symmetric utility specification, the problem becomes ordinal: triples can be rationalized based solely on the ordering of parties on a given dimension. The main results remain similar, with the exception that the Left–Right and the economy dimensions collapse into one because they imply the same party ordering.

To ensure comparability over time, I use CHES 2014 party positions throughout the sample. This separates changes in reported triples from changes in parties' positions, which are modest on the main dimensions over the period. Replicating the analysis with CHES 2019 scores yields the same conclusions (Figure A8). This is an important result: it suggests that the declining fit of the Left–Right dimension is not an artifact of changing party positions.

[12] These overlap patterns admit a natural interpretation. Consider a society in which every individual votes ideologically along a single dimension, but the relevant dimension differs across individuals: some voters evaluate parties with the Left-Right dimension, others with the anti-elite dimension.

7.2 Test of important dimensions of party competition

The previous section compares the shares s_d across dimensions and over time. A natural question is whether these shares are informative about the ideological lines of voters, or whether uninformative dimensions would do just as well on the same data. The test in this subsection supplies that benchmark by comparing the observed s_d with the distribution of rationalizable shares generated by random dimensions.

To run this test, I simulate 100,000 draws d' of party positions from $U[0, 10]$, compute the implied distribution of $s_{d'}$ for each study year, and compare the observed s_d across years and real dimensions d to these values. The end of this subsection also provides a robustness exercise in which the distribution of $s_{d'}$ is computed by reshuffling the party positions; this exercise provides a harder test, as it uses real party configurations.

Take the economy dimension as an example. The null hypothesis is H_0: *Voters vote ideologically, but not along the economic dimension* Y_d, and the alternative is H_1: *Voters vote ideologically along the economic dimension* Y_d. Ideological voting is maintained under both hypotheses: the test contrasts one candidate ideological space with an uninformative one, and is not a test of spatial voting against unstructured choice. I treat rejection of H_0 as evidence that *voters are ideological along a given dimension*.

Figure 4 shows the 90%, 95% and 99% quantile values of $s_{d'}$. Hypothesis testing can be carried out by comparing the observed values of rationalizability to the upper quantiles (95%, 99%) of rationalizability for given years. The Left–Right and the economy dimensions outperform the 95% quantile in all years; the anti-elite dimension equals or outperforms the 95% quantile from 2018 onward (with the exception of the 2020 COVID-19 pandemic year) and reaches the 99% quantile value in 2022. Other dimensions do not surpass these upper quantiles in any year.

Table 1 presents more detailed information – test p-values, $\mathbb{P}(s_{d'} \geq s_d)$, for all dimensions and years. The p-values suggest that the shares of rationalizable triples along the Left-Right and economic dimensions are unlikely to be generated by random party positions in any year; the same holds for the anti-elite dimension in the later years.

Figure A9 constructs a more conservative benchmark by reshuffling party positions across parties for a given dimension. This corrects for the potential over-permissiveness of the test that uses uniformly distributed hypothetical party positions. The results of this analysis remain quantitatively similar.

7.3 Left-Right dimension is the best possible fit

As is clear from Table 1, among the 100,000 randomly constructed dimensions the Left-Right dimension rationalizes the largest share of observed triples from 2017 onwards ($\mathbb{P}(s_{d'} \geq s_d) = 0.0000$) and nearly the largest share in 2013–2016. This exercise, however, does not allow us to claim that the Left-Right dimension provides the best possible fit among *all* single dimensions. I now turn to this exercise.

Table 1: $\mathbb{P}(s_{d'} \geq s_d)$

Year	Left–Right	Economy	Anti-elite	Environment	Religion	Culture	EU
2013	**0.0004**	**0.002**	0.22	0.22	0.43	0.65	0.12
2014	**0.0004**	**0.002**	0.25	0.22	0.46	0.57	0.10
2015	**0.0004**	**0.002**	0.21	0.23	0.43	0.58	0.09
2016	**0.0004**	**0.002**	0.17	0.29	0.43	0.66	0.09
2017	**0.0000**	**0.002**	0.11	0.25	0.41	0.64	0.13
2018	**0.0000**	**0.002**	**0.05**	0.29	0.46	0.63	0.10
2019	**0.0000**	**0.010**	**0.05**	0.30	0.50	0.72	0.11
2020	**0.0000**	**0.010**	0.15	0.30	0.47	0.73	0.07
2021	**0.0000**	**0.003**	**0.02**	0.20	0.40	0.62	0.21
2022	**0.0000**	**0.010**	**0.01**	0.27	0.40	0.60	0.25

Notes: The table shows the probabilities that the share of triples rationalized by a random dimension, $s_{d'}$, is at least as large as the share rationalized by a given dimension, s_d. The distribution of $s_{d'}$ is constructed by randomly sampling party positions from the interval $[0, 10]$ (this is the interval for the CHES data) 100,000 times; 0.0000 means no generated dimension performed better. s_d are calculated based on the CHES 2014 party positions. Values below or equal to the significance level $\alpha = 0.05$ are in bold. *Note: Data collection in 2020 differed due to the COVID-19 pandemic (Glasenapp et al., 2024).*

To do this, I construct for each year the best possible dimension. While this exercise might seem difficult at first, note that under the symmetric utility the set of triples a dimension rationalizes depends only on the ordering of the parties and on the midpoints between party positions: perturbations of positions that preserve these orderings leave the rationalizability pattern unchanged. The search for the best dimension—the dimension whose set of jointly rationalizable triples has the largest observed frequency—therefore reduces to a finite problem, which I solve by enumerating all 360 distinct party orderings (up to reflection) and searching over the rationalizability patterns within each ordering. By the same logic, infinitely many dimensions attain the maximal share: any sufficiently small perturbation of party positions leaves the set of rationalizable triples, and hence the rationalized share, unchanged. The question is therefore whether the CHES Left-Right dimension belongs to this set of best dimensions.

The first row of Table 2 reports the answer. The results mirror Table 1: from 2017 onwards the Left-Right dimension rationalizes the largest possible share of triples. In 2013–2016 it falls short of the best possible dimension by less than half a percentage point. This result allows me to conclude that in later years *the Left–Right dimension is the best possible way to model voters' behavior in a one-dimensional voting space.*

Table A4 reports the results of an exercise in which I relax the assumption that utility is symmetric around the bliss point. Once any single-peaked utility is allowed, the exact party positions no longer matter and a dimension is characterized solely by the order of parties it induces; I therefore ask whether the order induced by the Left-Right dimension is the best among all 360 distinct orderings (up to reflection). The results remain similar to those presented above: in most years the Left-Right order is the best possible

Table 2: The Left-Right dimension across the two exercises

	2013	2014	2015	2016	2017	2018	2019	2020	2021	2022
L-R is the best dimension	No	No	No	No	Yes	Yes	Yes	Yes	Yes	Yes
	(-0.02)	(-0.42)	(-0.23)	(-0.22)						

Notes: The Table reports whether the CHES Left-Right dimension rationalizes the largest share of observed triples attainable by any single dimension in the given year. The best possible dimension is found by enumerating all distinct party orderings and searching, within each ordering, over party positions consistent with that ordering. For years in which the Left-Right dimension is not the best, the difference between the largest possible rationalizable share and the Left-Right share is reported in parentheses, in percentage points; the best-performing dimensions in those years are the variants of Left-Right, with the AfD placed at the far left next to the Left.

order, and in the remaining years it is the second best (2014–2016) or third best (2019).

This result parallels the main focus of Apesteguia and Ballester (2025), who propose to find the best possible order of survey questions given the respondents' answers.[13] Figure A10 also adds the anti-elite dimension to the picture, demonstrating that the main results from the previous analysis hold – the overall rank of the anti-elite dimension order among all possible party orders is decreasing over time.

7.4 Non-rationalizable triples

The Supplemental Appendix presents the results on the shares of non-rationalizable triples. Non-rationalizable triples are defined as those triples that cannot be rationalized by any of the dimensions considered in the main text. The share is about 20% in the years until 2019, and then declines steadily, reaching 14% by 2022 (Figure A12). Taken together with the gradual decline in the share of triples rationalizable under the best-fitting dimension (Left–Right), this suggests that some voters may be shifting attention to other, less traditional, dimensions.

I also provide two heterogeneity results. First, the share of observed non-rationalizable triples is consistently higher in the East, especially in earlier years (2013–2017) – Figure A13. The result supports existing evidence that voters in post-communist countries still need to adapt to democratic norms and rules (see, e.g., Spenkuch, 2018 on how individuals in Eastern Germany learned how to cast strategic votes). It may also reflect a relative lack of information about parties, underdeveloped ideological preferences, or simply voting behavior that is inconsistent with ideological voting along any of the considered dimensions. Moreover, in Eastern Germany, respondents older than 45 choose non-rationalizable triples about 2 percentage points more often than younger respondents, suggesting a role for generational experience.

Second, the share of observed non-rationalizable triples is monotonic in the reported level of government approval. Those respondents who hold more negative views about the current government are more

[13] Note that in their setup the utility function is not restricted to be symmetric, and, hence, their main object of interest is the best order. The result here is stronger, because it highlights the importance not only of the order, but also of the distances between the parties.

likely to choose non-rationalizable triples – Figure A14, consistent with the idea that those individuals might be less involved in democratic processes.

Note that the results in the previous sections point to a declining fit over time: the share of triples rationalizable along the best-fitting dimension (Left-Right) falls, and so does the metrics of the attainable fit of any single dimension (see, e.g., the 99% quantile values in Figure 4, or Figure A10 for rationalizable shares across party orderings). The results in this section, however, show that the share of triples that cannot be rationalized along *any* of the considered dimensions also declines. Taken together, these patterns suggest the growing importance of non-traditional ideological dimensions and increasing disagreement within the electorate about which dimension matters most.

7.5 Partial party rankings and perceived important problems

So far, the analysis uses only respondents' top three parties and does not use any additional information about their preferences. A natural question is how this measure of rationalizability relates to voters' answers to other questions. To explore this, I study answers to the questions *"What is the most important problem in Germany nowadays?"*, *"What is the second most important problem in Germany nowadays?"*, which are asked in the same GLES survey, and correlate them with reported partial rankings.

If the partial-ranking test captures dimensions of competition that are actually important to voters, this should be reflected in their survey answers: respondents who cite economic issues as most important should be more likely to have triples rationalizable along the economic dimension than respondents who do not cite economic issues. Note that this comparison does not depend on the overall salience of any given issue; it only requires discernible differences in issue reporting across voters with different triples.

These questions are open-ended, and the survey vendor codes responses into broad categories (159 in total). I manually grouped these categories into four groups:

1. *Economic problems* (concerns about economic situation, social benefits, unemployment and taxes),
2. *Government-related problems* (mentions of coalition formation, political disillusionment, disputes within the government, outcome of the elections and rising right-wing extremism),
3. *Climate change* (mentions of environmental protection, climate change and renewable energy),
4. *EU problems* (mentions of Eurozone crisis, EU free trade agreement with the USA and others).

The corresponding dimensions are economy, anti-elite, environment, and EU, respectively[14]. I next correlate the chosen answers with triple rationalizability for a respondent i, separately for each dimension

[14] Only dimensions that rationalize on average more than 10% of observed triples are considered. I drop the Left–Right dimension because I cannot identify a coherent set of "most important problems" that cleanly maps into it.

d:

$$\text{Rationalizability}_{id} = \beta_0 + \beta_1 \, \mathbb{1}[\text{Topic}_{id}] + \epsilon_{id}, \tag{1}$$

where $\text{Rationalizability}_{id}$ is an indicator equal to one if respondent i's triple is rationalizable along dimension d, and $\mathbb{1}[\text{Topic}_{id}]$ equals one if respondent i mentions an issue associated with dimension d among the most important problems. Table 3 reports the estimated β_0 and β_1 from specifications for different d.

Table 3: Relationship between triple rationalizability and the frequency of corresponding self-reported most important problems, by dimension

	Rationalizability along:			
	Economy	Environment	Anti-elite	EU
	(1)	(2)	(3)	(4)
Economic problems	0.024***			
	(0.002)			
Climate Change		0.022***		
		(0.002)		
Government related problems			0.016***	
			(0.003)	
EU problems				0.070***
				(0.005)
Intercept	0.421***	0.156***	0.230***	0.386***
	(0.002)	(0.001)	(0.001)	(0.001)
Observations	171,079	171,079	171,079	171,079
Adjusted R^2	0.001	0.001	0.0002	0.001

Notes: *p<0.1; **p<0.05; ***p<0.01. Only dimensions that rationalize on average more than 10% of observed triples are considered. The estimates are β_0, β_1 from the Specification 1. *Economic problems* (concerns about economic situation, social benefits, unemployment and taxes), *Climate change* (mentions of environmental protection, climate change and renewable energy), *Government-related problems* (mentions of coalition formation, political disillusionment, disputes within the government, outcome of the elections and rising right-wing extremism), and *EU problems* (mentions of Eurozone crisis, EU free trade agreement with the USA and others).

The results indicate that voters who prioritize a given set of problems are more likely to choose triples consistent with those concerns. Effect sizes relative to the means are modest, especially for economic issues. This likely reflects the overall salience of different issues, as opposed to their importance for voting choices.

Taken together, these findings suggest that the rationalizability measure is not mechanical: it aligns with voters' stated concerns and appears to capture underlying motivations for party choice. Relative to self-reported views or priorities, this approach more directly reflects how voters trade off issues at the moment of choice and how these trade-offs translate into party selections.

7.6 Misreporting

An important limitation of questions about respondents' preferred party, and other survey questions prone to socially desirable responding, is their vulnerability to social image concerns (Bursztyn et al., 2020; DellaVigna et al., 2016). These concerns are especially salient for stigmatized parties at the political extremes. In Germany, this primarily manifests as underreporting of support for the AfD (Valentim, 2021). I show that this concern is less severe for tests of ideological voting than for other exercises, such as estimating ideal points of voters.

Misreporting of a stigmatized party occurs when individuals report a lower utility from supporting it than they actually experience. This can manifest in two ways: either by excluding the party entirely (placing it last in the ranking) or by understating its utility (ranking it lower than it would otherwise be).

While these two forms are difficult to disentangle empirically, the survey data provide clear evidence of misreporting. Table A2 compares the share of respondents naming each party in response to "Which of these parties do you like best?" with the parties' Bundestag vote shares in 2013, 2017, and 2021. Consistent with prior evidence, the survey share is systematically lower for the AfD in all years. The second form of misreporting, ranking the AfD lower rather than omitting it, appears quantitatively limited: the share of respondents listing the AfD as their second preference is 1.8%, 2.5%, and 2.0% in 2013, 2017, and 2021, respectively (and 2.8%, 2.0%, and 1.9% as their third preference). These magnitudes are too small to close the gap with vote shares, let alone to account for voters who would truthfully rank AfD second and third.

However, if an individual has ideological preferences along dimension d, she can still rank all parties in the choice set and thus provide a well-defined partial ranking. Moreover, when the stigmatized party is extreme, the misreported partial ranking still allows a researcher to detect ideological voting along dimension d under a single-peaked, not necessarily symmetric, utility function. Since the AfD is extreme on each policy dimension among the parties considered (Table A1), and the one-dimensional results are similar under symmetric and non-symmetric utility, I conclude that the one-dimensional findings are robust to misreporting. Additional details are provided in the Appendix B.

7.7 Two-dimensional ideological space

A natural extension is to run the same test in a multi-dimensional policy space, allowing individuals to place different weights on different dimensions. In practice, however, the available data do not provide enough variation for such an exercise to be informative.[15]

Recognizing the relevance of multidimensional voting, I conduct the following exercise: I identify the pair of dimensions, and the associated weights, that rationalizes the largest share of observed triples. In other words, I ask: *if all voters evaluate parties using the same two dimensions and the same relative weights, which dimensions and weights best explain the observed data?* I implement a grid search over weights on a 0.1 grid $(0/1, 0.1/0.9, \ldots, 1/0)$. As before, I benchmark these results against the 99th

[15]This result parallels the result in Apesteguia and Ballester (2025) on the lack of empirical power in a multidimensional space. The main difference between the two setups is that I assume a more restrictive utility function.

percentile of outcomes obtained from random dimensions for a given set of weights[16]. Table 4 reports, for each year, the pairs of dimensions and associated weights that best rationalize the observed data.

Table 4: Best Dimension-Pair By Year (With 99% Quantile Benchmark)

Year	2013	2014	2015	2016	2017	2018	2019	2020	2021	2022
Dimension 1	Economy	Economy	Economy	Economy	Economy	Economy	Economy	Economy	Economy	Economy
Dimension 2	Culture	Culture	Culture	Culture	Culture	Culture	Culture	Culture	Culture	Culture*
Relative Weights	0.60/0.40	0.60/0.40	0.60/0.40	0.60/0.40	0.60/0.40	0.60/0.40	0.60/0.40	0.60/0.40	0.60/0.40	0.50/0.50
Rationalized Share	0.876	0.867	0.885	0.855	0.822	0.800	0.814	0.854	0.785	0.781
99% Quantile	0.811	0.782	0.802	0.770	0.739	0.705	0.743	0.780	0.718	0.704

* In 2022, the economy–culture pair with 0.60/0.40 weights attains the same rationalized share; there are no exact ties in other years.
Notes. The table reports, for each year, the highest share of rationalizable triples ("Rationalized Share") obtained from two-dimensional models. Each model is defined over a given pair of dimensions ("Dimension 1" and "Dimension 2") and a set of relative weights searched over a grid (0/1, 0.1/0.9, . . . , 1/0) ("Relative Weights"). The reported value corresponds to the dimensions-weight combination that yields the highest rationalized share in a given year. The 99% Quantile values are constructed from the distribution of rationalized shares obtained using pairs of random dimensions, each independently uniformly drawn from [0,10] (the support of the CHES data), across 5,000 simulations. Quantiles are computed separately for each weight combination, since different weights can, in principle, generate different rationalizable shares. Rationalized shares are the fraction of observed triples for which there exists a point in the 2D policy space such that, under quadratic distance-loss utility with weights (w_1, w_2; $w_1 + w_2 = 1$), the chosen first/second/third parties are respectively the closest options. Losses across dimensions are additively separable (no interaction term), so the total utility is the weighted sum of the two dimension-specific distance components.

The results point to a stable pair of the most relevant dimensions: economy and culture, with economy carrying the larger weight (60%) in all years except 2022. In 2022, the 60%/40% and 50%/50% weights yield similar shares of rationalizable triples, suggesting an increase in the relative importance of the cultural dimension. This suggests that while Left–Right constitutes the best one-dimensional model of voters' behavior, a two-dimensional model benefits from decomposing it into its core components (economy and culture) to obtain the best-fitting voting space. Table A3 reports the second-best combinations of dimensions and weights over the years.

These findings do not contradict the one-dimensional evidence on the rising importance of the anti-elite dimension; instead, they point to growing heterogeneity in which dimensions voters consider most salient. First, the maximum share of rationalizable triples achieved by the best-fitting two-dimensional specification declines over the sample period (Table 4), indicating that even the best two-dimensional model fits the data increasingly poorly—consistent with additional dimensions or non-ideological factors shaping choices. Second, Appendix Figure A15 reports the normalized rank (lower is better) of the Left–Right/anti-elite pair across weights: specifications that assign greater weight to anti-elite improve over time and, by 2021–2022, rank within roughly the top 7% of all dimension–weight combinations[17]. These patterns suggest that voters increasingly prioritize different ideological dimensions, making it harder to fit their choices with one- or two-dimensional models.

Unfortunately, the available data do not allow me to extend the analysis to three dimensions, which would permit a clearer assessment of the role of the anti-elite dimension and other factors behind the declining fit of the best-fitting models.

[16] Shares of rationalizable triples in a two-dimensional space are sensitive to weights, with equal weights across dimensions being on average the most permissive set of weights.

[17] The fine weight grid generates many near-equivalent specifications. Hence, the top 7% rank likely understates the role of anti-elite by counting many almost-identical pairs separately.

8 Discussion

Explaining why people vote the way they do is difficult, yet it is crucial for interpreting contemporary political landscapes and for informing forecasts and policy responses. This paper adopts a spatial voting framework and shows how survey-based partial party rankings can be used to identify ideological lines of voters.

Using the German Longitudinal Election Study (2013–2022) and CHES party placements, I find that the Left–Right dimension provides the best-fitting one-dimensional model of voting choices among all the possible dimensions in most years, and rationalizes about 40-50% of respondents' reported partial rankings depending on the year. In two dimensions, the best-fitting specification combines the economic and cultural dimensions, consistent with interpretations of the dominant Left–Right axis as bundling these components. However, the fit of both models deteriorates over time.

The declining performance of the traditional one- and two-dimensional models defined by the Left-Right and Economic-Cultural dimensions suggests that voters increasingly rely on other considerations when evaluating parties. One explanation is the rising importance of the anti-elite dimension: its explanatory power increases over time in the one-dimensional analysis and improves markedly in the two-dimensional analysis when combined with Left–Right. This pattern is consistent with the growing prominence of anti-elite (populist) parties in Germany and beyond. Once such movements gain sufficient support, political conflict can shift toward questions about the legitimacy of the establishment and existing institutions (Guriev and Papaioannou, 2022).

Other explanations include a broader shift in voting criteria toward other ideological cleavages or factors such as valence, identity, representation, or geography (Kendall et al., 2015; Ford and Jennings, 2020; Kawai and Sunada, 2025).

The framework developed in this paper can be applied to other countries and survey instruments that elicit ranked choices, and to multi-category party evaluations such as feeling thermometers. Richer data would allow future work to assess additional dimensions directly and to better model heterogeneity across voters.

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A Tables and Figures

Table A1: Party positions along different dimensions, CHES 2014

Party	Left–Right	Economy	Culture	Anti-elite	EU	Environment	Religion
AfD	8.923	8.333	8.692	7.778	1.615	8.667	7.000
CDU	5.923	5.917	6.000	0.800	6.385	5.364	6.200
FDP	6.538	8.000	3.385	0.900	5.692	6.909	2.300
Greens	3.615	3.500	2.154	2.000	6.231	1.455	2.000
Linke	1.231	1.250	4.923	5.400	3.000	4.778	1.667
SPD	3.769	3.501	4.154	1.300	6.386	4.273	3.100

Notes. These numbers are estimated party positions on different dimensions from the Chapel Hill Expert Survey (Jolly et al., 2022). For the 2014 survey, 337 experts evaluated 268 parties in all EU countries. Each dimension ranges from 0 (left, liberal, pro-elite, etc.) to 10 (right, conservative, anti-elite, etc.). *Left–Right* refers to overall ideology; *economy* to positions on privatization, taxes, regulation, government spending, and the welfare state; *culture* to the GAL–TAN dimension (social and cultural values); *anti-elite* to the salience of anti-elite rhetoric; *EU* to attitudes toward European integration; *environment* to environmental sustainability; and *religion* to the role of religious principles in politics. The positions of SPD and Greens and of SPD and CDU on the economy and EU dimensions, respectively, are identical; I shift the positions of SPD by 0.001 on both dimensions to resolve those ties. This does not affect the analysis, as Figure A11 demonstrates.

Table A2: Sample first-choice shares vs. official list (second) vote shares (%)

	2013			2017			2021		
Party	Sample	List vote	Δ (S–L)	Sample	List vote	Δ (S–L)	Sample	List vote	Δ (S–L)
AfD	**2.46**	**4.70**	**−2.24**	**4.70**	**12.60**	**−7.90**	**4.62**	**10.40**	**−5.78**
CDU	39.54	34.10	5.44	34.72	26.80	7.92	22.73	19.00	3.73
CSU	5.80	7.40	−1.60	7.56	6.20	1.36	8.62	5.20	3.42
FDP	2.17	4.80	−2.63	7.53	10.70	−3.17	10.17	11.40	−1.23
Left	8.22	8.60	−0.38	11.21	9.20	2.01	7.00	4.90	2.10
SPD	23.47	25.70	−2.23	23.13	20.50	2.63	20.75	25.70	−4.95
Others	3.00	4.50	−1.50	0.21	1.80	−1.59	0.74	2.90	−2.16

Notes. Sample refers to survey answers to *"Which of these parties do you like best?"* among respondents who indicate any party. List vote refers to the official nationwide second-vote results in Bundestag elections. "Others" aggregates NPD, Pirates, and Free Voters (FW). Δ (S–L) is Sample minus List vote. Some of the differences between survey results and electoral results might seem contradictory. For instance, the FDP shows negative misreporting in 2017, even though it is usually not considered a stigmatized party and should not receive a lower share in the survey than in the election. Note, however, that the share is aligned with the polls conducted throughout 2017 – the FDP was expected to receive 6–7% throughout the first half of the year, but its share increased right before the elections and reached more than 10% (Politico, 2024). Hence, the share in the survey seems to be a weighted average of these polls. In contrast, the share of the AfD was volatile but never fell below 8%, and the respective shares do indicate misreporting.

Table A3: Second Best Dimension-Pair By Year (With 99% Quantile Benchmark)

Year	2013	2014	2015	2016	2017	2018	2019	2020	2021	2022
Dimension 1	Environment	Environment	Environment	Economy	Environment	Economy	Economy	Economy	Economy	Economy
Dimension 2	Economy	Economy	Left–Right	Culture	Left–Right	Culture	Culture	Culture	Culture	Culture
Relative Weights	0.60/0.40	0.60/0.40	0.60/0.40	0.70/0.30	0.60/0.40	0.70/0.30	0.70/0.30	0.70/0.30	0.70/0.30	0.60/0.40
Rationalized Share	0.865	0.852	0.876	0.841	0.813	0.789	0.807	0.843	0.779	0.781
99% Quantile	0.811	0.782	0.802	0.759	0.739	0.690	0.734	0.775	0.689	0.722
Diff (Best - Second)	0.011	0.015	0.009	0.014	0.009	0.011	0.007	0.011	0.006	0.000

Notes. The table reports, for each year, the second-highest share of rationalizable triples ("Rationalized Share") obtained from two-dimensional models. Each model is defined over a given pair of dimensions ("Dimension 1" and "Dimension 2") and a set of relative weights searched over a grid (0/1, 0.1/0.9, . . . , 1/0) ("Relative Weights"). The reported value corresponds to the dimensions-weight combination that yields the second-highest rationalized share in a given year. The 99% Quantile values are constructed from the distribution of rationalized shares obtained using pairs of random dimensions, each independently uniformly drawn from [0,10] (the support of the CHES data), across 5,000 simulations. Quantiles are computed separately for each weight combination, since different weights can, in principle, generate different rationalizable shares. Rationalized shares are the fraction of observed triples for which there exists a point in the 2D policy space such that, under quadratic distance-loss utility with weights (w_1, w_2; $w_1 + w_2 = 1$), the chosen first/second/third parties are respectively the closest options. Losses across dimensions are additively separable (no interaction term), so the total utility is the weighted sum of the two dimension-specific distance components.

Figure A1: Second-choice shares conditional on first choice (diversion ratios)

Notes: Each cell reports the share of respondents naming the row party as their first choice and the column party as their second choice, normalized by the first choice. Shares are aggregated across the ten survey years (GLES 2013–2022); cell shares in each row sum to 100%. Parties are ordered along the Left-Right dimension.

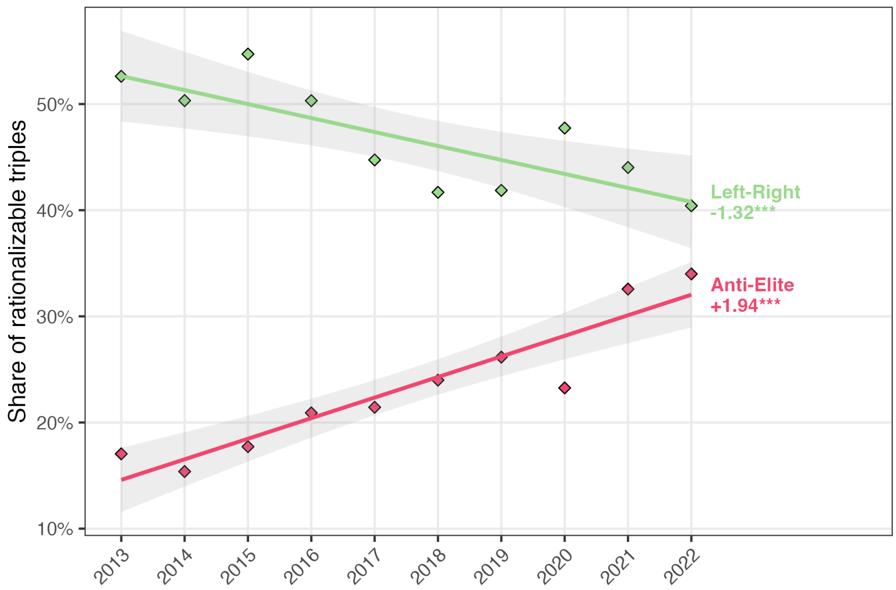

Figure A2: Linear trend of the shares of rationalizable observed triples across years for the Left-Right and anti-elite dimensions.

Notes: Diamonds show the yearly shares of rationalizable triples in one-dimensional models for the Left-Right and anti-elite dimensions. Party positions are fixed at CHES 2014 for all years. Solid lines are fitted linear trends with 95% confidence bands; the figures next to each line report the slope of the trend in percentage points per year, estimated from an individual-level linear probability model of the rationalizability indicator on a linear time trend; * p<0.1, ** p<0.05, *** p<0.01.

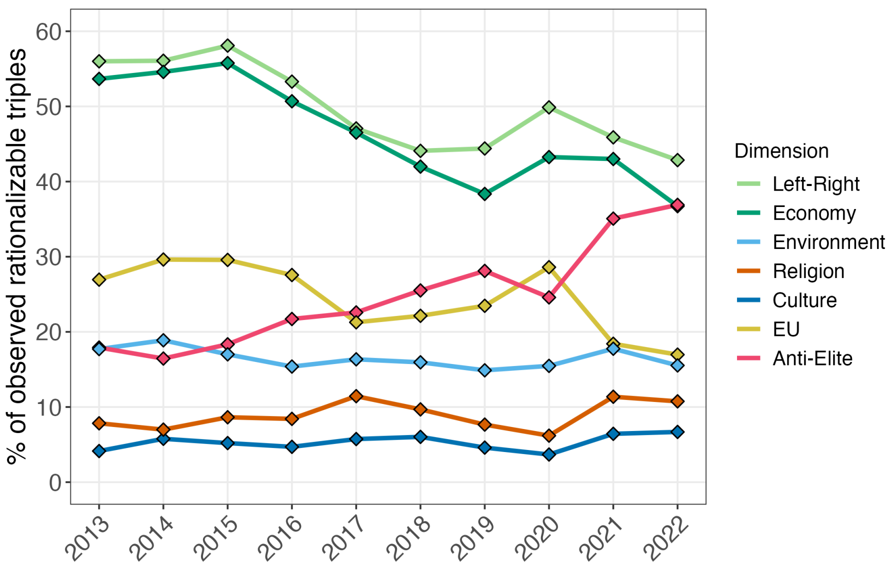

Figure A3: Share of rationalizable triples across years and dimensions, excluding the AfD from the analysis.

Notes: The sample is restricted to respondents who do not list AfD among their three preferred parties; AfD is also removed from the set of parties used to compute rationalizability. Triple shares are renormalized each year to sum to 100%, so the scale is comparable to the main results. The rise of the anti-elite dimension is visible in this subsample, indicating that it is not driven by AfD supporters alone.

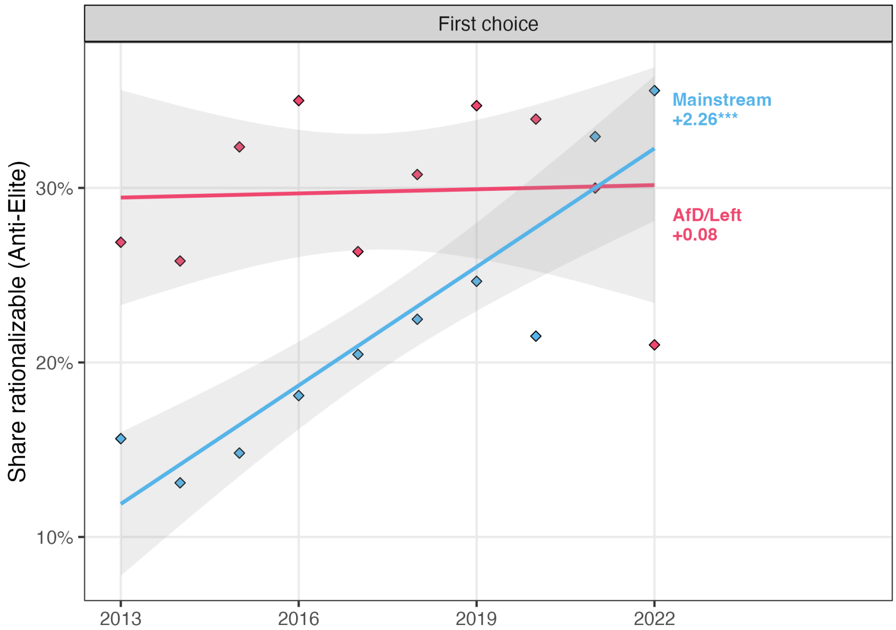

Figure A4: Share of rationalizable observed triples across years for the anti-elite dimension separately by supporters of mainstream and non-mainstream parties.

Notes: Respondents are split by their first-choice party: non-mainstream (AfD or Left) versus mainstream (CDU, SPD, FDP, Greens). Diamonds show the yearly share of respondents in each group whose observed triple is rationalizable along the anti-elite dimension (GLES 2013–2022; party positions from CHES 2014). Solid lines are linear trends fitted to the individual-level data, with 95% confidence bands. The figures next to each line report the estimated slope of the linear trend in percentage points per year; * $p<0.1$, ** $p<0.05$, *** $p<0.01$.

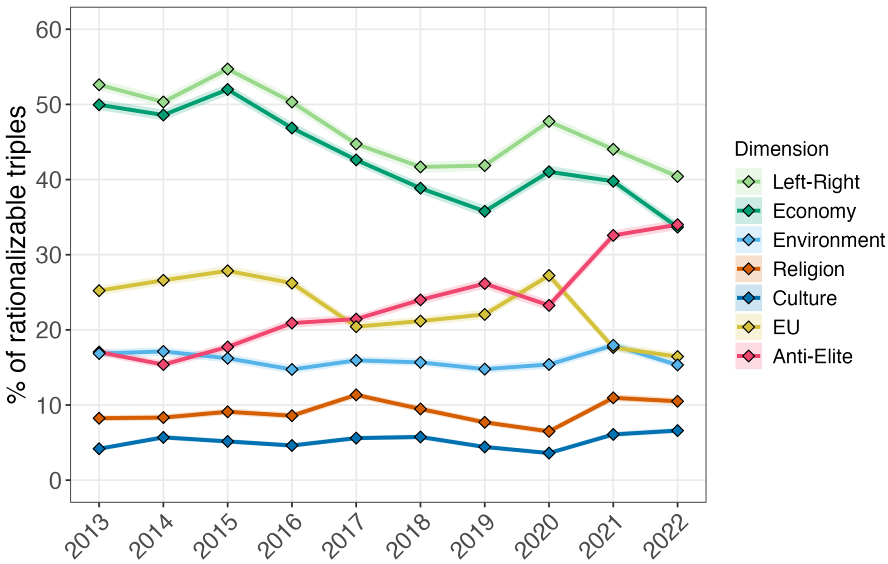

Figure A5: Share of rationalizable observed triples across years and dimensions.

Notes: Each line shows the share of respondents whose observed triple is rationalizable along the given ideological dimension in a one-dimensional spatial model, by survey year (GLES 2013–2022). Party positions are taken from CHES 2014. Shaded bands are 95% bootstrap confidence intervals: within each year, respondents are resampled with replacement 1,000 times, holding the observed yearly number of respondents fixed and giving each respondent equal probability of being drawn; the bands connect the 2.5th and 97.5th percentiles of the resampled shares.

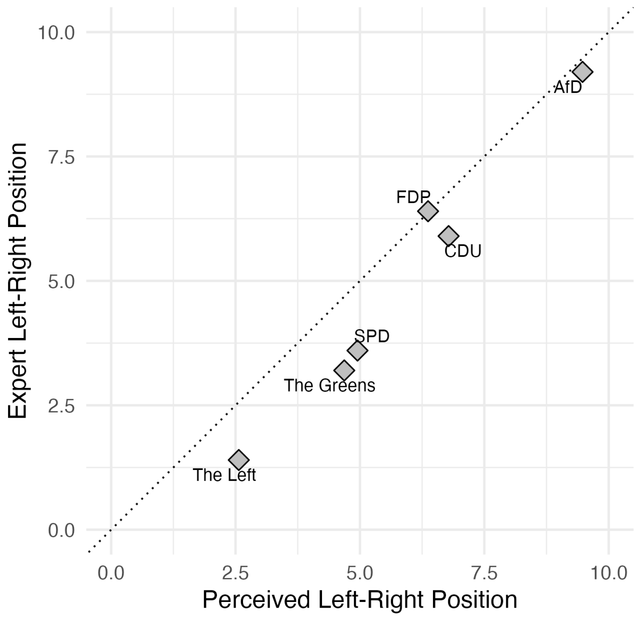

Figure A6: Perceived party positions (by respondents) and CHES positions on the Left–Right dimension

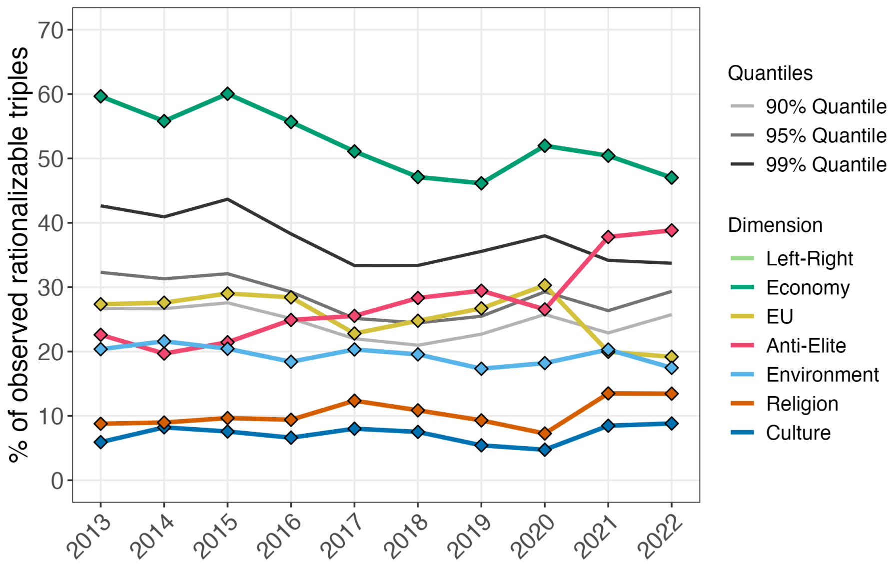

Figure A7: Share of rationalizable observed triples across years and dimensions allowing for non-symmetric utility functions around the bliss point.

Notes: The plotted values are the shares of rationalizable triples for one-dimensional models for different years allowing for non-symmetric utility functions around the bliss point. Gray lines denote 99%, 95%, 90% quantiles of the share of rationalizable triples based on 100,000 simulations of random party positions.

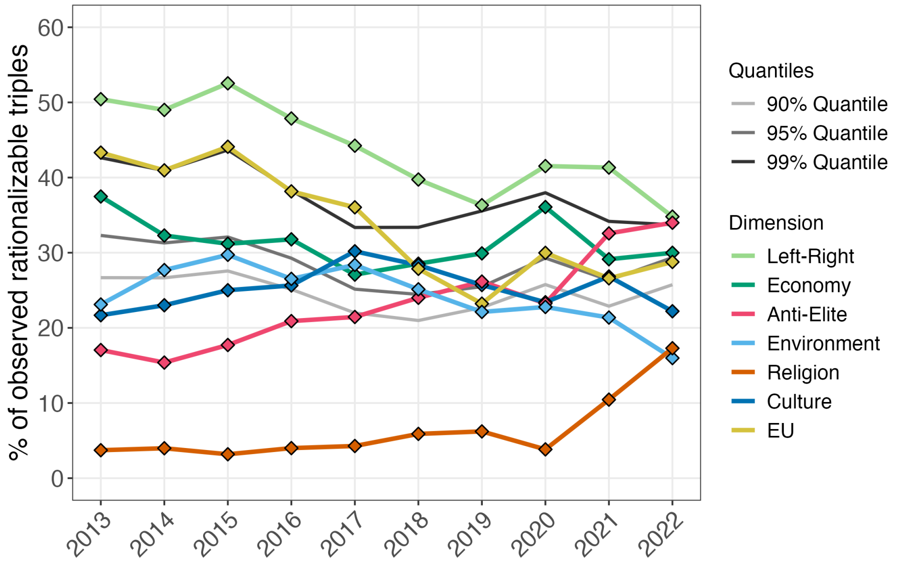

Figure A8: Share of rationalizable observed triples across years and dimensions.

Notes: The plotted values are the shares of rationalizable triples for one-dimensional models for different years using the CHES 2019 party positions. Gray lines denote the 99%, 95%, and 90% quantiles of the share of rationalizable triples based on 100,000 simulations of random party positions.

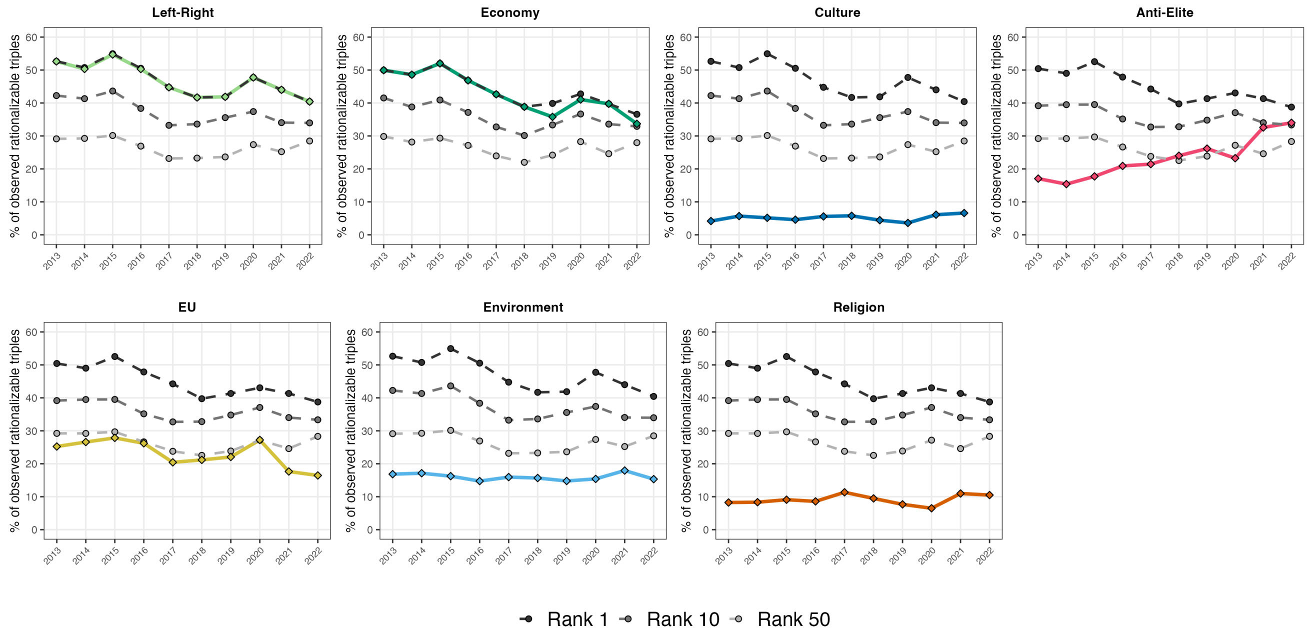

Figure A9: Share of rationalizable observed triples across years and dimensions, with rank values from the permutation test.

Notes: Each panel plots the share of rationalizable triples for a given one-dimensional model across years (solid colored line). Rank values are computed by permuting party labels across the six observed party positions within each dimension: rank 1 means the true labeling achieves the highest rationalizability share among all 720 permutations, rank 10 that it falls within the top 10, and so on. The permutation test benchmarks the true ordering against all empirically plausible alternatives, preserving the distribution of positions.

Table A4: The Rank of the Left-Right dimension order among all possible orders

	2013	2014	2015	2016	2017	2018	2019	2020	2021	2022
Rank of the Left-Right order	1	2	2	2	1	1	3	1	1	1

Notes. The Table reports the rank of the Left-Right *ordering* among all 360 orderings by the share of triples rationalizable with each ordering (1 = best); in this exercise a triple is counted as consistent with an ordering if it can be rationalized by some single-peaked preference (not necessarily symmetric) over that ordering.

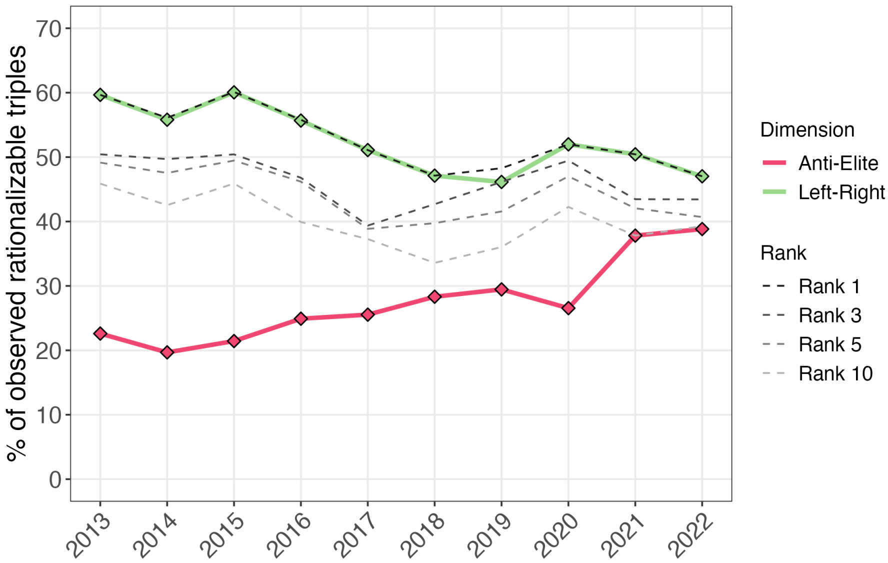

Figure A10: Share of rationalizable triples under **ordinal** distances for the Left-Right and anti-elite dimensions, with ranks over the universe of all possible one-dimensional party orderings.

Notes: The plotted shares are computed under ordinal (rather than cardinal) party distances. The rank is based on all $6!/2 = 360$ distinct one-dimensional orderings of the six parties (the factor of $1/2$ reflects that mirror-image orderings yield identical rationalizability sets). Rank 1 means the Left-Right (or Anti-Elite) ordering rationalizes a larger share of observed triples than any alternative one-dimensional ordering in that year.

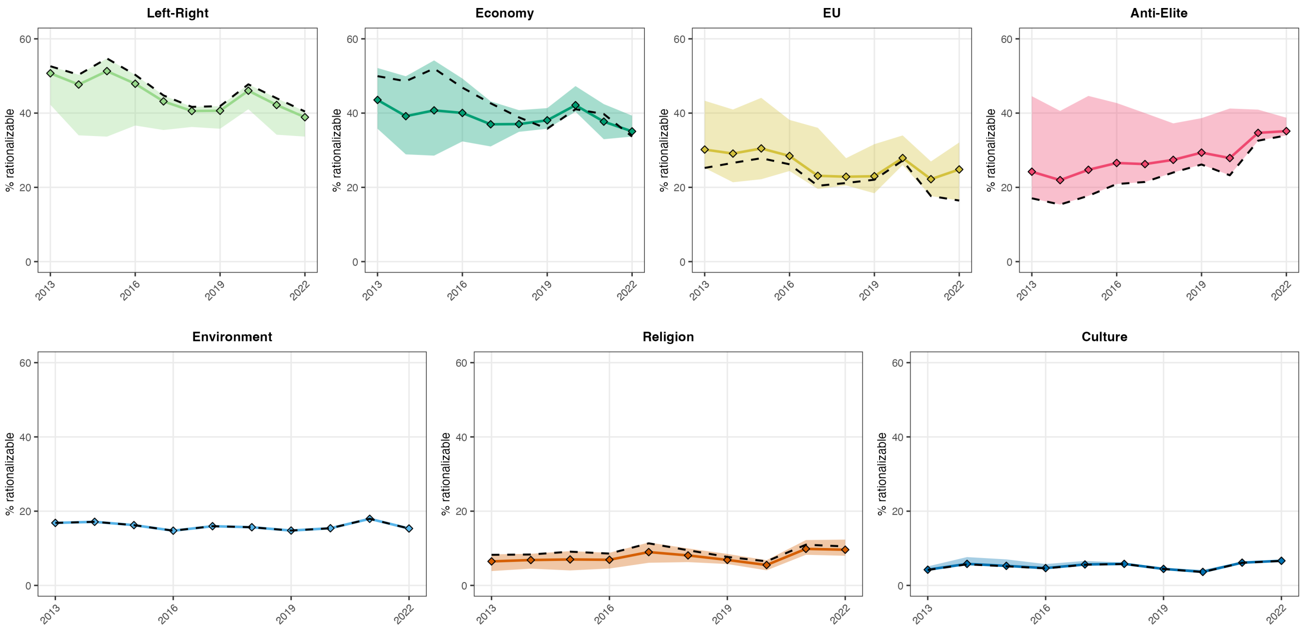

Figure A11: Robustness of the main results to noise in party positions: 1,000 draws from a normal distribution centered at CHES estimates with standard deviation 0.1.

Notes: For each dimension, 1,000 sets of party positions are drawn from independent normal distributions centered at the CHES 2014 estimates, with standard deviation 0.1. For each draw, rationalizability shares are computed. Black dashed lines show rationalizability at the true CHES positions; colored lines show the mean across simulations; shaded bands are empirical 95% confidence intervals. The estimates at the true positions lie within the confidence bands in every year, confirming robustness of the results to small and uncorrelated measurement/perception errors.

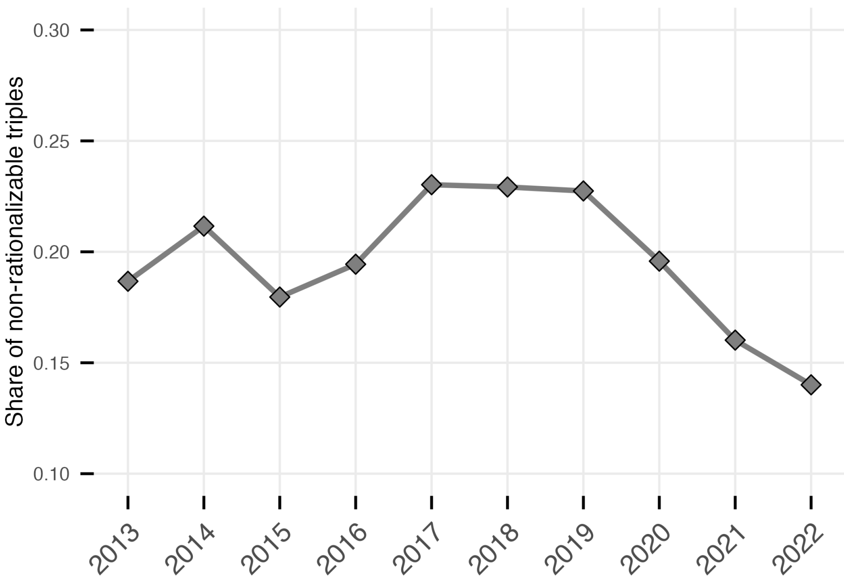

Figure A12: Share of non-rationalizable observed triples across years.

Notes: The plotted values are the shares of non-rationalizable triples. A non-rationalizable triple is defined as a triple that cannot be rationalized along any of the considered dimensions: Left–Right, economy, anti-elite, environment, religion, culture, EU.

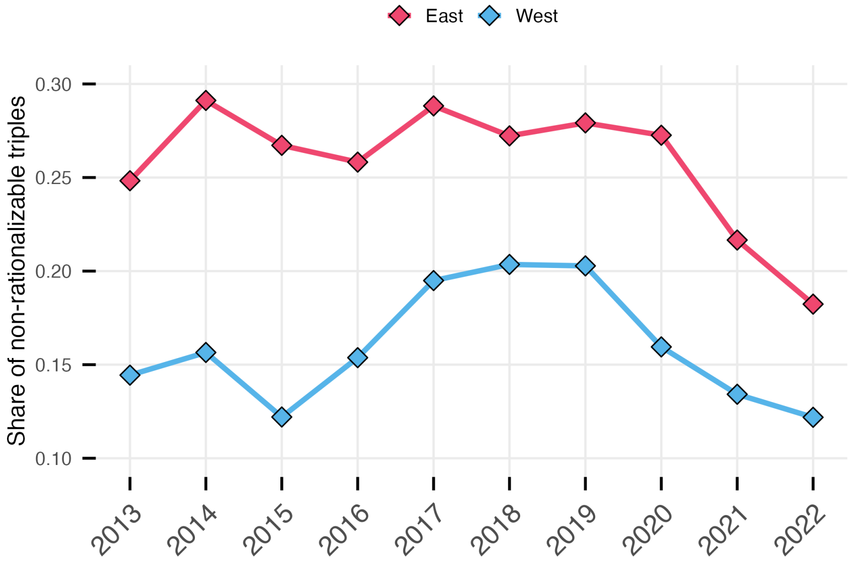

Figure A13: Share of non-rationalizable observed triples across years and East and West Germany.

Notes: The plotted values are the shares of non-rationalizable triples. A non-rationalizable triple is defined as a triple that cannot be rationalized along any of the considered dimensions: Left–Right, economy, anti-elite, environment, religion, culture, EU. East refers to individuals from Eastern Germany, West to individuals from Western Germany.

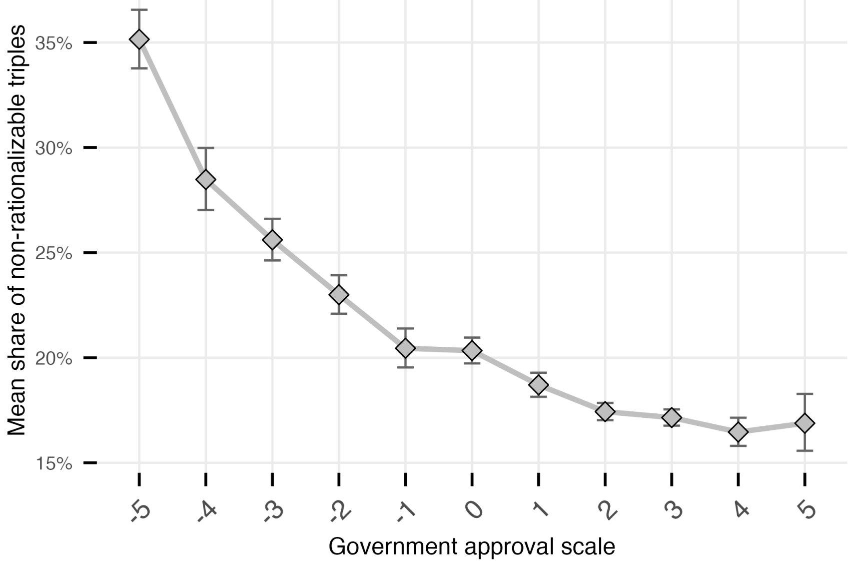

Figure A14: Share of non-rationalizable observed triples by reported level of government approval.

Notes: The plotted values are the shares of non-rationalizable triples by the levels of government support. A non-rationalizable triple is defined as a triple that cannot be rationalized along any of the considered dimensions: Left–Right, economy, anti-elite, environment, religion, culture, EU.

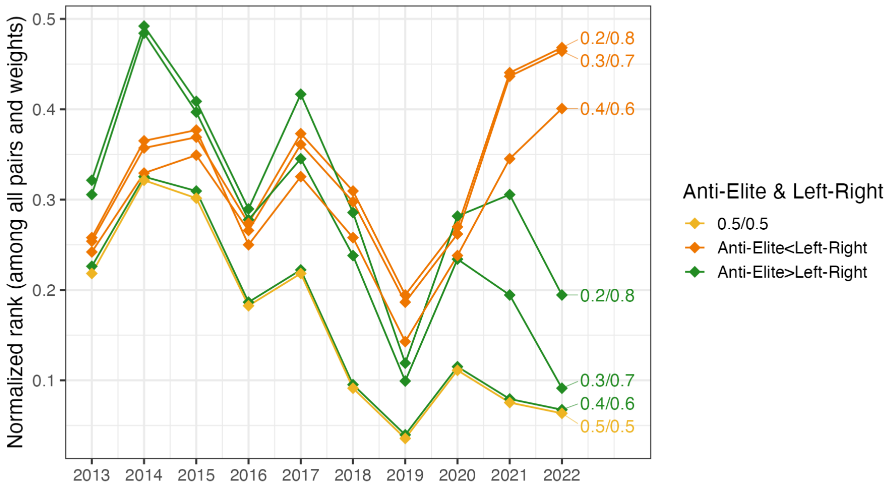

Figure A15: Normalized rank of the Left–Right/anti-elite pair for different weight combinations and years.

Notes: The plotted values are the normalized ranks (absolute rank divided by the total number of specifications) for the Left–Right/anti-elite two-dimensional model across years and weight combinations, where rankings are based on the share of rationalizable triples. Labels to the right of each line report the corresponding weights, with the first number denoting the weight on anti-elite and the second the weight on Left–Right. Rationalized shares are the fraction of observed triples for which there exists a point in the 2D policy space such that, under quadratic distance-loss utility with weights (w_1, w_2; $w_1 + w_2 = 1$), the chosen first/second/third parties are respectively the closest options. Losses across dimensions are additively separable (no interaction term), so the total utility is the weighted sum of the two dimension-specific distance components.

B Misreporting: Additional Discussion

As noted in the main text, an important concern is misreporting of preferences for the AfD. This may take two forms: respondents may (i) omit the party entirely (effectively placing it at the bottom of the ranking), or (ii) understate its utility by ranking it lower than they otherwise would. In this section I show that, in a one-dimensional voting space, such behavior does not affect rationalizability under the assumption of a single-peaked (not necessarily symmetric) utility function.

The exercise is as follows. Consider a truthful ranking that includes an extreme (stigmatized) party and is rationalizable in a one-dimensional policy space. I argue that the corresponding reported ranking—which is identical except that it either omits the stigmatized party or places it lower—is also rationalizable.

Consider for simplicity 4 parties a, b, c, d, where a is stigmatized and located at an extreme of the one-dimensional policy space. Suppose the voter votes ideologically along this dimension, i.e. her preferences are single-peaked around a bliss point (not necessarily symmetrically). In one dimension, a ranking is rationalizable if and only if it is consistent with moving away from the bliss point along the line; equivalently, it can be characterized by the neighbor structure of party positions.

Suppose the truthful ranking is $a \succ b \succ c \succ d$. Since a is extreme, it has a unique neighbor on the line. If a is the voter's top choice in a single-peaked order, then the second-ranked party must be this unique neighbor, i.e. b. Party b has two neighbors, and the truthful ranking implies that the interior neighbor of b is c. With four parties, it follows that c and d are neighbors. Hence, the reported ranking obtained by omitting a, namely $b \succ c \succ d$, is itself a contiguous single-peaked order and therefore rationalizable. The case in which a is not ranked first but is omitted is analogous.

Similarly, if an individual reports the stigmatized party a lower in the ranking, the reported order can still be rationalized by some bliss point. Intuitively, because a is extreme, shifting the bliss point toward the interior makes a farther away relative to the interior parties, while preserving a single-peaked ordering.

Symmetric utility. Robustness to misreporting under symmetric utility is ultimately an empirical question, because it depends on the configuration of party positions (in particular, the distances between parties). I illustrate this in Figure A16. Under the party configuration in panel **A**, misreporting can lead to a non-rationalizable report even when the truthful ranking is rationalizable. For example, if the truthful ranking is $a \succ b \succ c \succ d$, then the report that omits party a, $b \succ c \succ d$, is not rationalizable (in the presence of a). Intuitively, this is because party d is farther from b and c than party a is. By contrast, under the configuration in panel **B**, both the truthful ranking $a \succ b \succ c \succ d$ and the reported ranking $b \succ c \succ d$ are rationalizable.

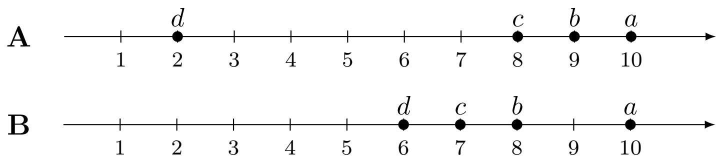

Figure A16: Policy lines with hypothetical party configurations a, b, c, d

Since the one-dimensional results under non-symmetric utility closely mirror the main results, I interpret the one-dimensional conclusions as robust to misreporting. There is no analogous general robustness result in two dimensions.

Robustness to misreporting under symmetric utility for one dimension is a property of the specific party configuration, and I check it directly for the two dimensions carrying the main one-dimensional results. For the Left-Right and anti-elite dimensions the results are robust to misreporting: for every truthful triple involving the AfD that is rationalizable under symmetric utility, the corresponding triple with the AfD omitted or demoted remains rationalizable. Hence the symmetric one-dimensional shares s_d for these dimensions are, empirically, unaffected by AfD misreporting.

C Computing Rationalizable Shares

Consider a reported triple $v = (a, b, c)$: party a is chosen from the full party set J, with $q = |J|$, party b from $J \setminus a$, and party c from $J \setminus a, b$. Under ideological voting for given d with a bliss point located at y in Y_d, the profile v is chosen if and only if

$$\mathrm{dist}(y, y_d^a) < \mathrm{dist}(y, y_d^b) < \mathrm{dist}(y, y_d^c) < \mathrm{dist}(y, y_d^k) \quad \text{for all } k \in J \setminus a, b, c, \tag{2}$$

a system of $q-1$ strict inequalities: all remaining pairwise comparisons follow from transitivity of the distance ranking. A triple is rationalizable if and only if *some* $y \in Y_d$ satisfies (2). Under utility symmetric around the bliss point in one dimension, the distances are absolute differences, $\mathrm{dist}(y, y_d^j) = |y - y_d^j|$; each inequality in (2) then restricts y to an open half-line bounded by the midpoint of two party positions, so the solution set is an interval, and the triple is rationalizable if and only if this interval is non-empty.Under single-peaked (not necessarily symmetric) utility, preferences are no longer represented by a common distance function, and (2) is replaced by an ordinal condition: v is rationalizable if and only if each successive choice is adjacent to the previous one in the order of parties along d, once earlier choices are removed — equivalently, if and only if a, b and a, b, c consist of consecutive parties in that order. In two dimensions with weighted quadratic utility, each inequality in (2) is linear in y because the quadratic terms cancel, so rationalizability is a linear-programming feasibility check too. In each case, s_d is computed exactly; simulation enters only through the construction of benchmark party positions.