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DOI
10.1057/s41295-024-00383-7

Publication date
2024

Document Version
Final published version

Published in
Comparative European Politics

Comparative European Politics (2024) 22:839–864
https://doi.org/10.1057/s41295-024-00383-7

ORIGINAL ARTICLE

Can we aggregate voters' perceptions of political parties' left–right positions? Formal and probabilistic tests of the left–right scale as a unidimensional common space on cross-national and longitudinal data

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Accepted: 3 May 2024 / Published online: 1 July 2024

Abstract
The left–right scale is widely assumed to be a common space, a joint yardstick that facilitates political communication. Aggregate voter perceptions of party positions on the left–right scale are widely used by scholars, a.o. to test models of voting behaviour, assess voter-elite congruence, or study party system change. Remarkably, while these models hinge on the longstanding assumption that voters have a joint understanding of the ordering of political parties on the left–right scale, this assumption has not been put to a systematic test. This paper introduces formal and a probabilistic tests of the formal demands of a common space: individual transitivity and collective transitivity. Cross-national analyses of election survey data (36 countries in the CSES) and longitudinal analyses in Germany (1983–2021), Great Britain (1997–2019), and the Netherlands (1981–2021) test whether the left–right dimension meets these demands. The outcomes are sobering. They cast serious doubt on the interpretation of the left–right scale as a common scale among voters, except under specific circumstances. We discuss the far-reaching implications of these findings.

Keywords Left–right dimension · Voters · Common space · Spatial model · Transitive ordering

Introduction

The left–right scale is part of the elementary toolkit in electoral research. Election studies frequently ask respondents to position political parties on a left–right scale (e.g. CSES 2023). Electoral scholars have aggregated these individual perceptions to a collective ordering of parties from left to right, for instance to test models of voting behaviour (e.g. Best and McDonald 2011; Merrill and Grofman 1999), to measure the congruence between voters, parties, and governments (e.g.

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Blais and Bodet 2006; Golder and Stramski 2011), to assess voters' responsiveness to shifts in party programme, leadership, or message (e.g. Adams et al. 2014; Wagner and Meyer 2023), or for various other reasons (e.g. Dalton and McAllister 2015; Aldrich et al. 2018; Imre 2023; Hansen and Pedersen 2014).

The widespread scholarly use of aggregated perceptions of parties' left–right positions are based on the premise that the left–right dimension is a so-called common space (e.g. Jou and Dalton 2017; Klingemann 1995; Macdonald et al. 2007). The assumption of that common space requires voters to have a shared (basic) understanding of not only the left–right dimension but also of parties' positions on that dimension. This assumption has a long scholarly history. In 1974 Sani argued that the left–right scale is 'a socially shared mental construct that is used to pattern and give meaning to that aspect of reality that is ordinarily called political', and therefore 'suitable for general communication exchanges'. It would function as 'a common yardstick in the description of political parties' (Van der Eijk and Niemöller 1983), and as 'the most useful currency for information exchange between voters and parties' (McDonald and Budge, 2005: 32).

The idea of the left–right dimension of party positions as a common frame of reference to voters hinges on the assumption that 'people can reasonably recognise the parties' positions on the Left–Right scale' (Jou and Dalton 2017). The intersubjective understanding is key to substantive applications of aggregated positions. Macdonald, Rabinowitz, and Listhaug (1997) explicitly formulate as 'a cardinal principle of both the Downs model and the directional model that voters have a common view of the political space'. Building directly on this assumption, voters' aggregated perceived left–right positions of political parties are both widely used empirically and defended theoretically as meaningful. These aggregated perceptions would be these parties' *true* position (Macdonald et al. 2007: 412). Yet, it is not at all evident that this assumption of a unidimensional common left–right dimension of party positions holds (Gilljam 1997).

Evidently, the left–right scale has been put under a lot of scrutiny. Various studies aimed to inductively understand how voters explain the terms left and right (e.g. Bauer et al. 2017; Lindqvist and Dornschneider-Elkink 2023). Yet, they did not deductively test the existence of the common space of party positions. Other studies illustrated how information costs and biases explain variation in perceived party positions (e.g. Dahlberg, 2013; Vegetti and Širinić, 2019), studied the role of measurement (e.g. Miwa 2018; Lo et al. 2014), or tested the unidimensional fit of the voters' perceptions of parties' left–right positions (e.g. Carroll and Kubo 2021). Yet, even these studies build on the assumption of the existence of a common space (Armstrong II et al., 2014: 41; Kurella 2017: 72). The assumption of a unidimensional common space of party positions among voters thus remained implicit and unchecked.

The unidimensional common space of party positions among voters should not be a theoretical assumption. It is ultimately an empirical question. We aim to test systematically to what extent the left–right scale meets the demands of a unidimensional common space among voters, so that scholars can aggregate voters' perceptions to measure party positions intersubjectively.

It is important to specify this aim. First, the fundamental concern is whether the left–right scale of perceived party positions has a *shared* meaning among voters. The common space hinges on two demands (Arrow 1963; Black 1958; Downs 1957): individual transitivity (voters should be able to position political parties on a single line), and collective transitivity (voters should be able to do position parties intersubjectively, i.e. in the same order). To test both we rely on formal analyses and nonparametric Mokken scale analyses of ordinal data. Second, this paper focuses on the intersubjective ordering of parties. Therefore, we do not focus on (errors to) their absolute position (cf., Van der Eijk 2001) or on the use of the scale for strictly subjective, ideosyncratic measurement (cf., Westholm 1997; Blais et al. 2001). Third, a shared understanding of party positions on the left–right scale does not require a substantive interpretation of the scale. This means we need not define the meaning of left and right, but 'merely' test to what extent the aggregate ordering is shared. Fourth, we explicitly focus on voters and their perceptions. This means we cannot make any inferences about the existence of the theoretical construct itself (which may be derived from other sources such as manifesto data), or about the existence of a common space among experts or politicians.

To the extent that the left–right scale of party positions does not meet the twin demands of individual and collective transitivity, the use of aggregated perceptions of party positions breaks down. It is imperative to assess to what extent the left–right scale adequately satisfies the twin demands, as it is not at all evident that it does (Gilljam 1997). In some countries the labels of left and right are hardly used (e.g. the United States where the terms liberal and conservative prevail); in others the multidimensionality of the political space complicates an unidimensional understanding of left and right (Carroll and Kubo 2021). To systematically test the left–right scale as a unidimensional common space, we present formal analyses (that do not model measurement errors) and probabilistic IRT models (that have a model of measurement errors). Employing election survey data, we comprehensively perform cross-national analyses across 36 countries and longitudinal analyses (1981–2021) in three democracies (Great Britain, Germany and the Netherlands) with different levels of party system fragmentation, openness, and polarization analyses across time. These comparisons allow us to assess the empirical validity of the common left–right scale under a range of conditions that invoke varying information costs on voters and that complicate an intersubjective ordering of parties.

Our results indicate that the assumption of individual transitivity is met in most countries while the assumption of collective transitivity is not. Voters do not share an understanding of parties' relative position on the left–right scale. The prevalence of idiosyncratic perceptions impairs the use of aggregated left–right scores of party positions in population survey data. While these perceptions might be highly informative in explaining individual voting behaviour, they are not very useful in constructing measures of party positions.

Left–right orientations among voters: idiosyncratic perceptions or common scale?

Left–right orientation

Left–right orientations would be an ideological shortcut for voters who have incomplete information of the implications of the different party policies for their own utility-income. The possibility of constructing this single dimension rests on the assumption that political questions are reduced to their bearing upon one crucial issue. Modern political science regards the left–right dimension as some sort of 'super issue' that clusters together economic, social, and ethical issues (Laver and Budge 1992). The left–right scale is then the dominant conflict dimension on which the political battle is fought (Van der Eijk and Niemöller 1983), the space of competition on which political conflicts will be fought 'regardless of how many cleavage and/or identification dimensions exists' (Sani and Sartori 1983, p. 330).

Perceived party positions: individual scale or common scale

The use of the left–right dimension as 'a common yardstick in the description of political parties' to scholars, media, and voters (Van der Eijk and Niemöller 1983, p. 229; see also Sani 1974) hinges on the degree to which scholars, media, and voters indeed have a common understanding of its meaning. This shared interpretation is a fundamental requirement to treat the left–right dimension as a common political scale. Of course, this requirement does not apply to scholarly work that interprets left–right positions as strictly subjective and that is agnostic about the existence of a common space (e.g. Blais et al. 2001; Gilljam 1997; Kedar 2015; Pardos-Prado and Dinas 2010). As long as scholars accept voters' sometimes idiosyncratic interpretation of the left–right scale (cf. Bauer et al. 2017; Lindqvist & Dornschneider-Elkink 2023) and—by extension—their subjectively understood party positions on that scale, theories of electoral behaviour can be tested at the subjective level: 'Although voters may at times be mistaken about these locations, it is their personal beliefs (…) that will guide preference formation' (Westholm 1997, p. 870).

Yet, many electoral studies approach voters' understanding of the left–right scale as a common political space, an intersubjective ordering of political parties that has a collective meaning to the whole electorate.[1] Voter perceptions of party positions are key to the existence of a common political left–right scale.

Some scholars have interpreted the aggregated (average) perceived positions of political parties as evidence for a unidimensional common space, because they are rather stable over time (Kroh 2007; Dalton and McAllister 2015) and rather congruent with other measures (Jou and Dalton 2017). Macdonald et al. (2007, p. 412) write about 'the party's 'true' (mean perceived) position'. Dalton and McAllister (2015) state: 'the public's Left–Right placement of parties is virtually identical to the evidence from party experts or the party elites themselves. (…) This implies that citizens and political experts see a common Left–Right space, in which parties compete for people's votes'. In other words, the existence of a common political space (recognised as such collectively by citizens) is both measured and supported by aggregating individually perceived party positions to the electorate's mean perceptions. The simple aggregation bypasses the fundamental requirement for a common scale of party positions among voters: it holds only when citizens have a shared understanding of party positions. By its nature, the commonality of the left–right scale cannot be determined via aggregate data, but only via individual data.

Others studied the limitations of perceived party positions on the left–right dimension at the individual level. Acknowledging the relevance of information costs, scholars have studied the informational complexity induced by political parties themselves in for instance their rhetoric (Wagner and Meyer 2023), policy shifts (Dahlberg, 2009), coalition behaviour (Adams et al. 2016), and leadership changes (Somer-Topcu 2017). They also account for problems of perceptual biases, for instance due to partisanship (Dahlberg, 2013), social identity (Vegetti and Širinić, 2019), or misperceptions (Carroll and Kubo 2018). These findings offer an explanation for spatial stretch (i.e. a theory of distance). Yet, while these approaches offer explanations for voters' misperceptions, they do so given the assumed existence of a common (unidimensional) space. The question is how much variation is acceptable in order to still assume an underlying common understanding of party positions on the left–right scale.

Finally, some scholars tested the dimensional properties of perceived party positions.[2] Specifically, Aldrich-McKelvey scaling (Aldrich and McKelvey 1977; Hare et al. 2015) has been used to test the unidimensional properties of the

[1] Aggregated perceptions are used for a wide variety of reasons, a.o., to test models of voting behaviour (e.g. Fazekas and Méder 2011; McDonald and Rabinowitz 1997; McDonald et al. 2007; Merrill and Grofman 1999; Sani 1974), to measure voter-party, voter-government, or party-government congruence (a.o. Blais and Bodet 2006; Ferland 2017; Golder and Stramski 2011; Meyer and Strobl 2016), to assess voters' responsiveness to parties' programmatic shifts (e.g. Adams et al. 2014; Seeberg et al. 2017) or leadership changes (Fernandez-Vazquez and Somer-Topcu 2019; Somer-Topcu 2017) or other programmatic information (Wagner and Meyer 2023), to determine the degree of party system stability (e.g. Dalton and McAllister 2015), to draw comparisons to expert- or manifesto-based measures (e.g. Aldrich et al. 2018; Best and McDonald 2011; Imre 2023; Van der Brug 1999), to assess voters' degree of political sophistication (e.g. Hansen and Pedersen 2014), or to map the programmatic trajectory of (new) parties (e.g. Enyedi 2005).

[2] Additionally, there have been studies into the role of measurement choices, for instance of middle categories (Miwa 2018) and scale length (Lo et al. 2014).

left–right scale. The conclusions of Carroll and Kubo (2021) challenge the idea of the left–right scale of perceived party positions as a unidimensional common space: 'These findings indicate that conclusions derived from applications using left–right perception data—such as party competition and congruence measures—may be affected by the complexity of left–right perception. This suggests there is reason to be cautious when using left–right placement data as a single-dimensional concept and especially when comparing across different contexts'. Moreover, even Aldrich-McKelvey scaling explicitly builds on the assumption of a common space (Armstrong II et al., 2014: 41; Kurella 2017: 72).

Hence, while the aggregated left–right scale continues to be used as a unidimensional common space, the demands for a unidimensional common space have not been tested directly and deductively.

Formal demands for a common scale

From the onset, Downs (1957) had made the demands explicit. The very first assumption of his Downsian model reads: 'The political parties in any society can be ordered from left to right in a manner agreed upon by all voters' (Downs 1957, p. 115). In other words, if the left–right scale constitutes a collective political space, voters do not need to agree on the *absolute* position of political parties on that scale or on the absolute distance between these parties, but should agree on their *relative* positions: there must be a collective transitive ordering of parties.

The use of mean or median scores to position parties on a common scale thus assumes that the distribution of each party on the left–right scale may be located on a single dimension with a collectively transitive ordering. But this assumption may be false. Single-peaked preferences are the characteristics of individuals, not of the collective (Black 1958). Each individual single-peaked preferential ordering can be placed on a single dimension, if and only if individuals agree on the ordering of parties from one extreme to the other (Downs 1957, p. 116).

The general demand of a collective transitive ordering consists of two demands that are more precise. The first demand is that *individual* voters should be able to make a transitive ordering of political parties on the left–right scale. In more technical terms, individual voters need to satisfy Arrow's axioms I and II of the relation R (Arrow 1963): '*at least as left as*'. Axiom I states that for any pair of parties x and y, either x is more leftist to y or party y is more leftist to x or the two parties are equally leftist. Axiom II states that the left–right rating between the different pairs of alternatives is consistent: if party x is rated left or equal to party y, and party y is rated left or equal to party z, then party x is also rated left or equal to party z. If both axioms I and II are satisfied, relation R is a (weak) ordering relation (Arrow 1963, p. 14). The adjective weak means that the ordering relation includes the possibility of equal rating. The axioms of Arrow are formal requirements to guarantee that the alternatives (i.e. political parties) are placed on a single line. For the purpose of this study these two axioms boil down to a single demand: each individual must be able to give some left–right score to all political parties (cf. Dalton 2011, p. 112; Hansen and Pedersen 2014; Jou and Dalton 2017). Simply by giving numerical values to all political parties on a scale, survey respondents satisfy the condition of individual transitivity: by the design of survey questions this results in a transitive ranking of these parties on a straight line. When all parties are rated, the individual satisfies the Axioms of Arrow and meets the demand of individual transitivity.

The second demand is that voters come to a *collectively shared* ordering, so that their left–right orientations are not just idiosyncratic perceptions. Voters need to have a *shared* understanding of the ranking of political parties on the left–right scale. For that purpose, we compare all the party orderings that individuals make and assess to what extent they are identical. The most likely shared orderings are those according to average or median positions; surely, these are the orderings assumed in the empirical literature.[3] This ordering can be used as a benchmark to determine whether all individuals have an ordering relationship *R* '*at least as left as*' (or even relationship *P* '*to the left of*'[4]) that corresponds with the ordering according to average or median positions on the left–right scale.

To the extent that the electorate as a whole comes to collective transitive orderings and relationship *R* or *P* at the individual level, we can speak of a common scale. These demands are rather strict, but by necessity. In everyday use, the terms left and right may be 'convenient tags that can be associated with different social objects without the burden of having to be precise' (Sani 1974). However, in more scholarly use (i.e. to determine party positions mathematically and apply these positions in a meaningful way), validity and precision are prerequisites.

Hypotheses

As the left–right scale for perceived party positions continues to be widely used as a common scale in electoral research, our working hypothesis is that citizens are able to meet these demands in large numbers. Formally, all voters who are subjected to the common space in explanatory models of electoral behaviour (i.e. all voters) should meet these demands. However, that is highly unlikely. Therefore, we loosen the demands so that at least a majority of the electorate (50% + 1) should meet them in order to speak of a common scale.

Our hypotheses (in order of difficulty of the demands) read as follows:

H1 Citizens are able to make a transitive ordering of all political parties on the left–right scale. [individual scale].

H2 Citizens are able to make a transitive ordering in accordance with the collective ordering along relationship *R*. [individual scales match joint scale *R*].

[3] The use of median positions (e.g. Blais and Bodet 2006; Ferland 2017) is less common, but does not change the problem fundamentally.

[4] Relationship P is stricter than relationship R. To satisfy relationship *P*, all parties must have a different score. Moreover, the parties should be ordered in agreement to the collective ordering. In the most extreme case, an individual who places all parties on the same spot (A = B = C), would meet relationship R for any possible ordering, but not P for any possible ordering.

H3 Citizens are able to make a transitive ordering in accordance with the collective ordering along relationship *P*. [individual scales match joint scale *P*].

Data and methods

Data

To test these hypotheses in a comprehensive manner, we conduct empirical analyses in cross-national as well as longitudinal perspective. We test the assumption of the transitive ordering in 36 countries with parliamentary election surveys that included the joint module of the Comparative Study of Electoral Systems 2011–2016. The CSES includes two-party systems such as the United States as well as fragmented multi-party systems such as Slovakia and Norway. Subsequently, we employ longitudinal parliamentary election survey data from three European countries with different electoral systems and effective number of parties in parliament (Eff.N_s in 2017) (Gallagher 2023): the United Kingdom with 6 election surveys since 1997 (majoritarian system, Eff.N_s averages 2.4 (range: 2.1–2.6), Germany with 11 election surveys since 1983 (mixed system with 5% threshold, Eff.N_s averages 3.9 (range: 3.2–5.6)), and the Netherlands with 13 election surveys between 1981 and 2021 (proportional system in a single district with 0.67% threshold, Eff.N_s averages 5.3 (range: 3.5–8.5)).[5]

The sampling methods, response rates, and modes of data collection differ across countries and over time. While this might be an issue to the construction of a common political space on these survey data, it is not an issue for this paper: We test the internal validity of the assumed common spaces on the very same data sets used in scholarly work to construct these common spaces.

Measuring individual and collective transitivity

We base our analyses on a single question battery in the CSES 2011–2016 module that was phrased as follows: 'In politics people sometimes talk of left and right. Where would you place [party A] on a scale from 0 to 10, where 0 means the left and 10 means the right? Using the same scale, where would you place [party B-J]'. The same question was asked longitudinally in the United Kingdom (since 1997), Germany (since 1983), and the Netherlands (since 1981).[6] We incorporate all political parties covered by the CSES module that were represented in parliament before

[5] For the UK, we rely on the cross-sectional data at britishelectionstudy.com. For the Netherlands on the cross-sectional data at dpes.nl via the data station SSH. For Germany, we used the Kieler Wahlstudie, 1983–1990 (ZA1399,1681,1959), KAS Nachwahlstudie, 1994 (ZA4909), CSES 1998–2005 (ZA3073,4216,4559), and GLES, 2009–2021 (ZA5301,5701,6801,7701),.all available at gesis.org.

[6] The length of the scale varied slightly over time in the Netherlands from a 10-point scale (between 1981 and 1998, and in 2003) to the now common 11-point scale (in 2002, and since 2006). While this variation affects the longitudinal comparison, it does not affect the test of transitivity.

or after the elections, except regional parties that only participated in a geographically limited set of districts, such as the Scottish National Party and Plaid Cymru in Great Britain.

The question battery allows us to test individual transitivity (i.e. the percentage of respondents able to position all parties) and collective transitivity (i.e. the percentage of respondents that rank parties in agreement with the ordering of mean perceptions). If citizens are able to position all political parties on the left–right scale, they meet the two axioms of Arrow and make an *individual transitive ordering*. We operationalize this by measuring substantive missing values, i.e. respondents who report to be unable or unwilling to position all or at least one political party on the left–right dimension.[7] Respondents who ended the survey before the left–right perceptions are excluded from the analysis. In the CSES data this is not always clear. For countries where types of missing values are unclear, we therefore used neighbouring questions to distinguish between non-participation and item-specific missings.

If citizens additionally order political parties in congruence with the supposed collective ordering by respondents' average party positions, they also meet demands *R* and *P*. Ultimately, c*ollective transitivity* is assessed by the share of respondents that order political parties in line with the aggregate ordering by the mean position of the parties (see Appendix 1).[8]

Probabilistic tests: Mokken scale

Formal modelling demands all individuals to agree with the collective ordering. It lack a theory of errors: basically any transgression of the aggregated ordering would challenge demands *R* and *P*. There is no good criterion how many and which types of transgressions are allowed before rejecting the assumption of transitivity. However, we need to allow the existence of error in the measurement of the latent construct that is the common left–right scale, as not all errors are the same. Mixing up two directly neighbouring parties should be less problematic to *R* and *P* than mixing up parties that should be ordered far apart.[9]

For that purpose we extend our formal tests with Mokken scale analyses. Mokken scale analysis is the pre-eminent probabilistic test of unidimensional hierarchy (i.e. fixed ordering) for ordinal data in a nonparametric way (Van Schuur 2003). In comparison with other models employed in this field, Mokken scale analysis is nonparametric. It offers a direct probabilistic test of the second principle of the

[7] The test of the demands for a common space as the basis for average party positions does not rely on the nature of these missing values. An analysis of these missing values could be of substantive interest for other purposes.

[8] Results are not affected if we use the ordering according to (interpolated) median positions instead. In the three country studies, we also inductively compiled a list of unique orderings to calculate which of those most respondents share. This did not affect our conclusions.

[9] We performed additional analyses on the large parties in the CSES that do not consider errors between parties that in their average are very close together (≤ 0.5) as challenges to P and R. Results are decidedly mixed (see Appendix 3).

unidimensional common space (intersubjective agreement on the ordering of parties). The gist of the analysis is the idea that a person who positions (leftwing party) A on, for instance, position 4 should be more likely to position more rightwing parties B, C, and D on or to the right of that position. The H-value of scale homogeneity compares the observed errors in the hierarchical scale (such as incorrect sequences of pairs of parties as ordered by respondents) to the number of errors expected if there was no relationship whatsoever. A Mokken scale is acceptable if the H-value exceeds 0.3, and strong if the H-value exceeds 0.5.

In two ways the Mokken scale is an easier test of transitivity than the strict interpretation of the formal demands. First, the Mokken scale is only calculated for the set of respondents that came to an individual transitive ordering (i.e. that positioned all parties): respondents with missing party positions are not part of this test. Second, the Mokken scale allows for errors. However, in another way the Mokken scale is slightly more difficult to validate than the formal demands *R* and *P*. The Mokken scale for polytomous data uses the information of all item steps, so that it may matter for the scale fit whether respondent X erroneously positions rightwing party B merely one position to the left of leftwing party A (small error) or multiple positions to the left (larger error). Note, however, that this demand makes the Mokken scale analysis more rather than less similar to the common use of mean party placements to create a common scale.

Results: collective transitivity across the globe (2011–2016)

Collective transitivity across the globe

Figure 1 illustrates to what extent citizens in 36 countries were able to meet the two demands for collective transitivity of party orderings. We find that support for a common scale among voters is very weak, and that variation is rather large.

In most countries a majority of the people are able to position at least one party on the left–right scale (cf. Best and McDonald 2011). Thailand (36%) is the only exception. The share of respondents that meet the two axioms of Arrow (positioning all parties) ranges from 15% in Thailand to 91% in the Philippines. A complete, individual transitive ordering is reached by only a minority of the people also in Brazil (36% successful), Montenegro (31%), the Netherlands (37%), New Zealand (47%), and Romania (39%).

A collective transitive ordering is much more difficult. The share of respondents meeting demand *R* ranges from 1% (in the Netherlands) to 50% (in the United States). The share of respondents meeting demand *P* ranges from 0% (in 16 countries) to 43% (in the United States). Evidently, it is easier to come to collective transitivity if the number of parties is low (see Appendix 3), but the success rate is low and varied even in less fragmented party systems.

The country that seemingly best meets the demands of collective transitivity is the United States. Most people (86%) are able to assign a left–right position to both political parties, Democrats and Republicans. The US is also the country where most citizens meet demands *R* (50%) and *P* (43%): Democrats to the left or equal

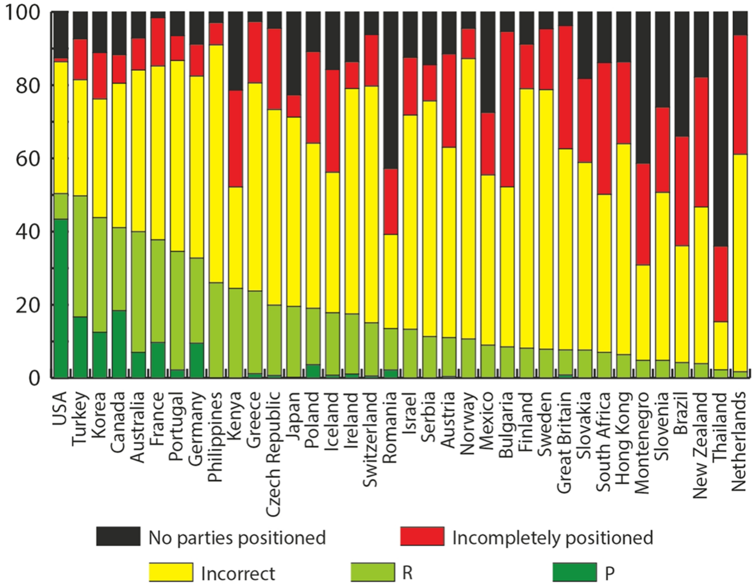


Fig. 1 Demands *P* and *R* (in percentages), CSES 2011–2017

to the Republicans. Yet, many Americans understand this position to be reversed: an only slightly smaller group of citizens positions the Republican party to the left of (36%) or equal to (43%) the Democratic party. In other words, even in this seemingly successful case—the only country where a majority of the citizens arguably meets the demand for collective transitivity—a very large minority of the citizens has a different interpretation of the underlying left–right scale (cf. Zaller 1992). That is remarkable, given that there are only three potential options ($D<R$, $R<D$, and $D=R$). While a vast majority of Americans comes to an individual transitive ordering, we cannot conclude that they have a shared understanding of the left–right scale.

All in all, these cross-national analyses reject the idea that voters have a common left–right scale for all parties that corresponds with the average placement on this scale. While we find support for hypothesis 1 in most countries, we reject hypothesis 2 in all countries except the United States, and reject hypothesis 3 in every single country under study.

Relaxing the demands

The formal demands of a full collective transitive ordering of political parties on the left–right scale are evidently very strict. Any transgression of the aggregated ordering counts as an error. Even though we allow half of the respondents to formulate a different or intransitive ordering, all countries except the United States fail to meet these demands. We may consider two reasons.

First, as we argued above (see paragraph 3.3) there is no theory of errors. Yet, even when we employ Mokken scale analysis as a systematic, probabilistic, and nonparametric test of unidimensionality of party positions among the subsample of respondents that assigned a position to all political parties in their country, we find no evidence for a shared understanding of relative party positions.[10] Only the Philippines ($H=0.38$), South Africa ($H=0.37$), and Montenegro ($H=0.32$) show left–right scales with a sufficiently homogenous fit (see Table 1).

Second, even though we only included political parties that participated nationwide, our test might contain too many parties that are irrelevant in practice. Consider, for instance, marginal parties such as the Greens in the United Kingdom in 2015 (3.8% of the votes, 0.2% of the seats) and the Orthodox Reformed SGP in the Netherlands in 2012 (2.1% of the votes, 2% of the seats). When we only include large parties in our analyses (minimally 10% vote share), the number of parties drops sharply and the success rate goes up (see Fig. 2). All countries have two to four parties above that threshold. A majority of the citizens meets demand R in 16 of the 36 countries, and demand *P* in 6. Yet, most countries continue to lack a shared understanding of the left–right scale by a majority of the electorate.

Overall, the assumption of a common left–right scale among voters does not find much support, except in specific countries and for a specific subset of the party system.

Results: three longitudinal case studies

Let us focus more in-depth on three countries with divergent electoral and party systems, for which longitudinal data on perceived left–right positions is available.

In the United Kingdom the share of respondents meeting demands *R* and *P* has dropped. Whereas in 1997 more than 50% of the respondents met demand *R*, it was met by no more than 10% in 2017, before it recovered to almost 30% in 2019 (see Fig. 3a). However, the recovery in 2019 coincides with a drop in the share of respondents meeting the demand for individual transitivity.

There are two tentative reasons why collective transitivity dropped between 1997 and 2017. First, the entrance of small, nation-wide political parties in parliament (the Greens and UKIP)[11] in 2015 and 2017 made it considerably more difficult to position all parties let alone meet demands *R* and *P*; by 2019 the Greens and the Brexit Party might have been better known. When we only focus on the three traditional parties—Labour, Tories, Liberal-Democrats—it is much less problematic to meet the demands to position all parties (see Fig. 3b). Second, the mean of particularly the Labour party shifted: the party of Tony Blair was perceived to be much more centrist in 2001 and 2005 than in earlier or subsequent waves, and practically overlapped with the average perception of the LibDem position. This

[10] The H-coefficients worsen when we disregard all irrelevant item steps and exclusively look at the sequence of items.

[11] We did not include parties that only participated in Scotland or Wales.

Table 1 Transitivity and scalability of perceived party positions, CSES 2011–2017

	Transitivity					Scalability (Mokken)
	None	Some	Incorrect	Order r	Order p	H-coefficient
Australia	7.1	8.7	44.2	33.0	7.0	−0.03
Austria	11.5	25.5	52.0	10.6	0.4	0.18
Brazil	34.0	29.9	31.9	4.2	0.0	0.28
Bulgaria	5.4	42.4	43.7	8.5	0.0	0.09
Canada	11.7	7.8	39.5	22.7	18.4	0.02
Czech Republic	4.5	22.2	53.5	19.2	0.7	0.06
Finland	8.8	12.2	70.9	8.1	0.0	0.18
France	1.6	13.2	47.5	28.1	9.7	0.08
Germany	8.9	8.7	49.7	23.3	9.5	0.20
Great Britain	3.7	33.7	55.0	6.8	0.8	0.10
Greece	2.6	16.8	56.9	22.6	1.2	0.15
Hong Kong	13.7	22.4	57.7	6.4	0.0	0.12
Iceland	15.7	28.1	38.4	17.1	0.8	0.16
Ireland	13.6	7.3	61.6	16.4	1.1	0.07
Israel	12.4	15.8	58.5	13.3	0.0	0.27
Japan	22.7	6.0	51.8	19.3	0.2	0.18
Kenya	21.4	26.5	27.7	24.5	0.0	0.10
Korea	11.1	12.7	32.4	31.3	12.5	−0.01
Mexico	27.5	17.0	46.5	8.9	0.1	0.21
Montenegro	41.4	27.8	26.1	4.8	0.0	0.32
The Netherlands	9.9	53.1	36.5	0.6	0.0	0.10
New Zealand	17.8	35.5	42.8	3.9	0.0	0.10
Norway	4.5	8.2	76.6	10.7	0.0	0.05
Philippines	2.9	6.1	65.0	25.9	0.1	0.38
Poland	10.9	25.0	45.1	15.4	3.6	0.00
Portugal	6.5	6.8	52.2	32.4	2.2	0.01
Romania	42.8	18.0	25.7	11.4	2.2	0.00
Serbia	14.4	9.9	64.4	11.3	0.0	0.17
Slovakia	18.2	22.9	51.3	7.6	0.0	0.22
Slovenia	26.0	23.3	45.9	4.8	0.0	0.09
South Africa	13.8	36.0	43.2	7.0	0.0	0.37
Sweden	4.6	16.7	70.9	7.8	0.0	0.08
Switzerland	6.2	14.1	64.7	14.5	0.6	0.09
Thailand	64.0	20.7	13.1	2.2	0.0	0.21
Turkey	7.4	11.2	31.7	33.1	16.7	−0.02
USA	12.6	1.1	36.0	7.0	43.4	−0.81

induced more errors in those years, not in the likelihood to position all parties but the likelihood to order these parties in line with their mean position.

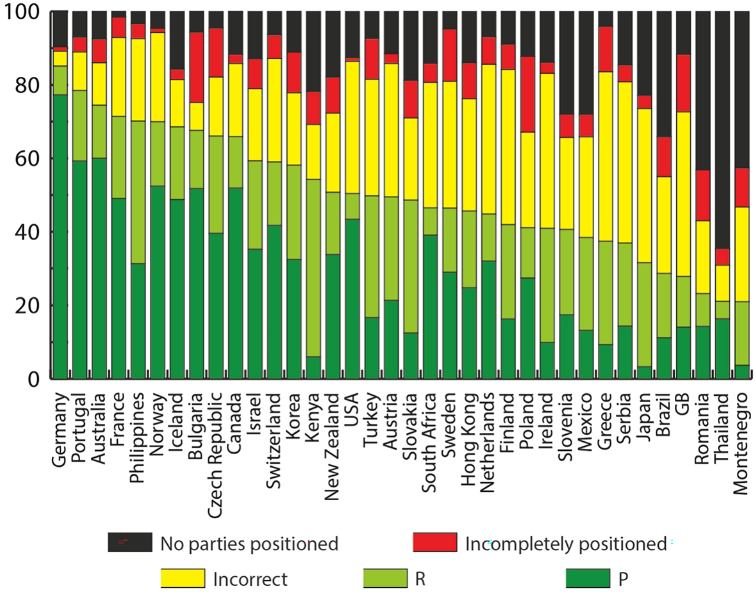


Fig. 2 Individual orderings of political parties with a more than 10% vote share, CSES 2011–2017

In Germany, too, respondents have become less likely to fulfil the assumptions behind the common left–right scale (see Fig. 4a). The share of respondents able to position all parties did not see much of a systematic or substantial decrease. Yet, the share of respondents meeting demands *R* and *P* declined sharply between a high point in 1987 (69.8% and 40.6%, respectively) to a low point in 2017 (35.1% and 6.6%, respectively). Most of the decline took place in the 1990s. There are multiple explanations. First, the German reunification led to the integration of East German voters into the (West) German party system. That contributed to the drop in *R* and *P*. West Germans showed higher *R* and *P* than East Germans, even though both would drop in the West from 1994 onwards. Second, the increase is partially due to the rise of new political parties on the left (e.g. die Linke, Greens) and right (e.g. AfD). However, even when we exclusively focus on the three parties that traditionally made up German coalition governments (SPD, FDP, CDU), a more structural decline of the commonality of the left–right scale can be discerned (see Fig. 4b). Third, ideological shifts and diversified governing coalitions at the state level signalled weakening and changing structures of party competition that likely made it more difficult for voters to position even these three parties correctly.

Finally, Dutch respondents are structurally much less likely to position all parties and to meet demands *R* and *P* than their British and German counterparts. This is, of course, connected to the number of parties in the Dutch party system. Although they fluctuate, these success rates show no trend in the Netherlands (see Fig. 5a). Moreover, we cannot relate these fluctuations to the sudden realignment of the Dutch party system in 2002 (cf. Pellikaan et al. 2007), to the total number of parties to position, or to the share of newcomers in the party system. When we focus

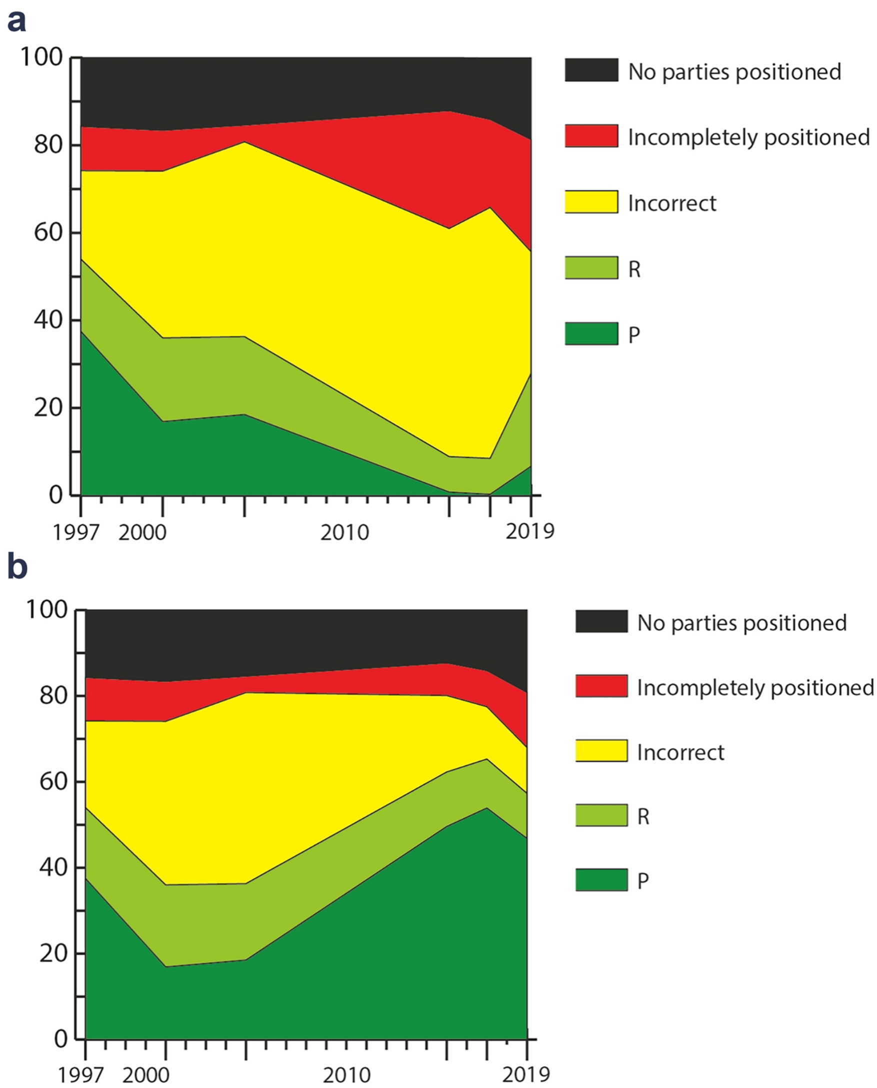


Fig. 3 **a** Demands *P* and *R* (in percentages), Great Britain 1997–2019 (all national parties) (The number of parties to be positioned changed over time from 3 (Tories, Labour, LibDem) in 1997–2005 to 5 in 2015 (adding Greens and UKIP)-2019 (Greens and Brexit Party)). **b** Demands *P* and *R* (in percentages), Great Britain 1997–2019 (three traditional parties)

exclusively on the four traditional government parties (PvdA, D66, CDA, VVD), similar trendless fluctuations are visible (see Fig. 5b).

These longitudinal analyses suggest that there is no structural increase or decline of the common space of perceived party positions. In the United Kingdom and Germany, support for the formal demands for a common scale responded to the rise of challenger parties, but in the Netherlands—traditionally much more familiar with may parties—it

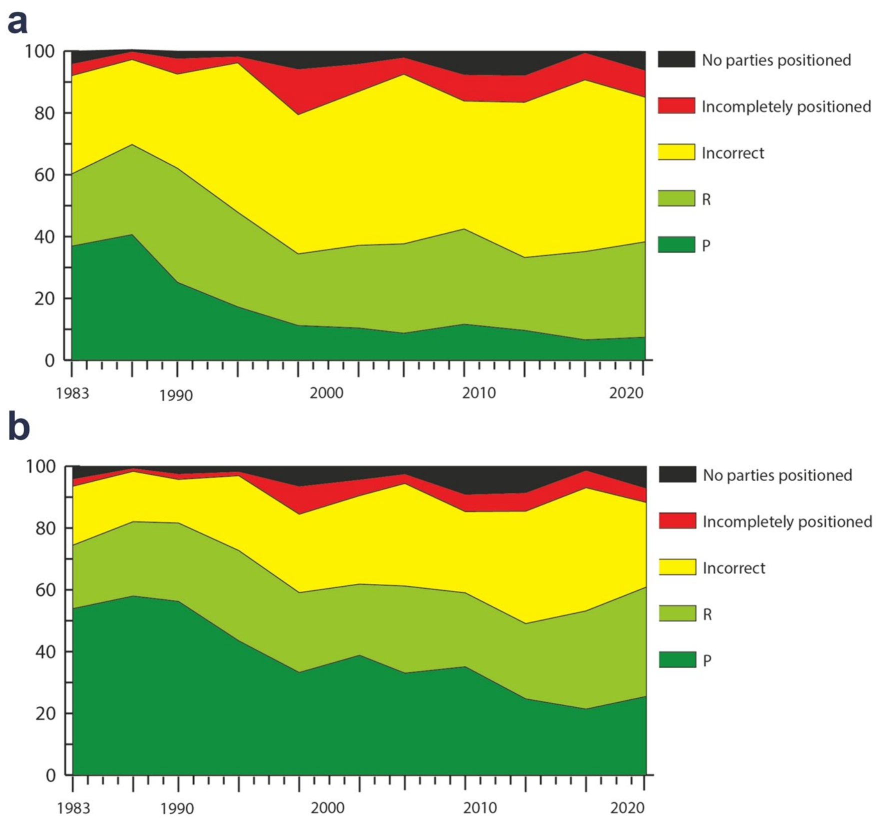


Fig. 4 **a** Demands *P* and *R* (in percentages), Germany 1983–2021 (all national parties). (The number of parties to be positioned changed over time from 4 (CDU, SPD, FDP, Greens) in 1983–1987 to 6 (adding PDS and Bündnis90) in 1990, to 5 (CDU, SPD, FDP, Greens and PDS/later Die Linke) in 1994–2013, and again 6 (adding AfD) in 2017 and 2021)). **b** Demands *P* and *R* (in percentages), Germany 1983–2021 (three traditional parties)

did not. When we exclusively focus on parties that traditionally made up government, we see an erosion of the commonality of the left–right scale only in Germany (primarily in the early 1990s), but not in the Netherlands or the United Kingdom. Rather, the United Kingdom underwent a temporary decline during the centrist government of New Labour. Rising numbers of (new) parties and depolarisation of the party system as a whole tend to complicate the common scale.

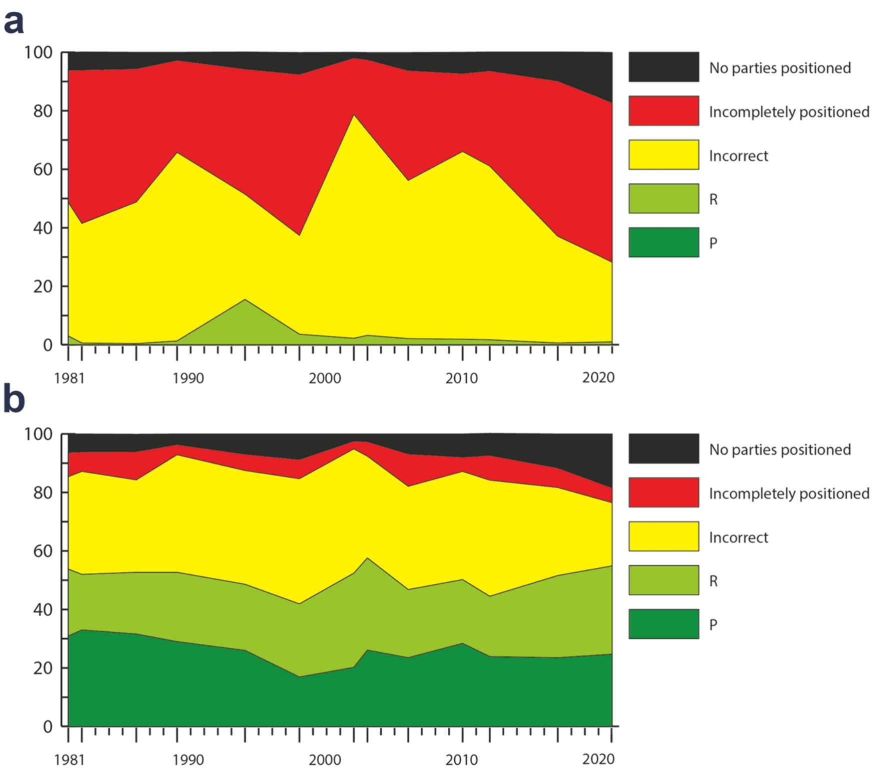


Fig. 5 **a** Demands *P* and *R* (in percentages), the Netherlands 1981–2021 (all national parties). (The number of parties to be positioned changed over time: 11 (1981), 13 (1982), 12 (1986), 9 (1989–1994), 11 (1998), 10 (2002–2010), 11 (2012), 13 (2017), 14 (2021)). Unelected parties are not included in the analysis. As the number of parties exceeded the number of available positions on the left–right scale, it is impossible to meet demand *P* in 1981, 1982, 1986, 1998, 2017, and 2021). (In 1981–1998 and 2003 voters were asked to position parties on a 10 point scale (1–10). In 2002, and 2006–2021 voters were asked to position parties on an 11-point scale (0–10)). **b** Demands *P* and *R* (in percentages), the Netherlands 1981–2021 (four traditional parties). (In 2002, and post-2006 voters were asked to position parties on an 11-point scale (0–10), instead of a 10-point scale (1–10) in other years)

Conclusion

Our findings question the assumption of the left–right dimension as a unidimensional common scale among the electorate. In the most favourable interpretation of our findings, most citizens are able to position all parties in their country on the left–right scale. They thereby meet the demand of individual transitivity, if only because the scholarly reliance on rating scales ensures an individually transitive ordering of parties. Yet, citizens do not have a shared (intersubjective) understanding of the ordering of political parties from left to right in their countries. The assumed common left–right scale is understood by a mere minority of the voters in nearly all countries, ranging from 1% in the Netherlands to 50% in the United States. Even when we allow for the existence of errors, IRT models show very little support for the existence of collective transitivity. The lack of conformity to a 'true' ordering of political parties is thus not merely indicative of measurement error, misperceptions or information costs suffered by voters (cf. Zaller 1992), but questions whether the left–right scale has an intersubjectively shared understanding altogether.

This conclusion is consequential. Theoretically, it directly challenges the 'cardinal principle (…) that voters have a common view of the political space' (Macdonald et al. 1997). Although the assumption of a unidimensional common space has been widespread (e.g. Downs 1957, p. 115; Jou and Dalton 2017; Macdonald and Rabinowitz 1997; Sani 1974; Van der Eijk and Niemöller 1983), it has not been put to systematic and rigorous test. Empirically, a diverse set of studies on party competition and electoral behaviour are founded on the assumption of the left–right scale as a common political space, most notably those that aggregate subjective perceptions of party positions to aggregated mean scores. Yet, because the left–right scale is not a common space, the aggregation of voters' perceptions is a problematic basis for analysis. Average perceptions of the positions of—and distances between—parties are not meaningful to individual voters who positioned these parties in the first place. Even when aggregate perceptions line up well with manifesto or expert positions (e.g. Jou and Dalton 2017; Van der Brug 1999), our findings question their theoretical value to explain the mechanisms behind individual voter behaviour. Aggregate perceptions are likely to induce misestimations of the theoretical model. Arguably, it could lead to underestimations as well as overestimations of relationships. The use of politically sophisticated voters' aggregate perceptions instead of those of the general electorate tends to offer no solution to this conundrum. Theoretically, the effects of these 'actual party positions' on individual voters need to be mediated by voters' individual, subjective perceptions, which in turn requires an individual level foundation to the assumption of a common scale.

All in all, we reject the assumption that the left–right scale is a one-dimensional common political space among voters, as there is no collectively intransitive ordering of political parties among the electorate. But what are the alternatives?

First, our findings do not reject the existence of the underlying construct of the left–right dimension in general. As we specifically tested its existence as a common space *among voters* and *on party positions*, we do not touch on the left–right scale as it is derived from other sources, such as content analysis of party manifestos, and expert opinions. The lack of evidence for a common left–right scale among voters does not mean it does not exist. Yet, this left–right scale cannot be assessed via the aggregation of voters' understanding of party positions.

Second, we do not contest the left–right dimension of party positions as a purely personal, subjective ordering. As a subjective rather than intersubjective (aggregated) measure, the left–right scale remains a valuable tool in the methodological toolbox of electoral research. Yet, subjective voter perceptions are not a simple solution, if only because they make tests of rivalling models of voting behaviour more sensitive to endogeniety. Fortunately, test outcomes are rather robust against the use of aggregated or individual orderings of parties (Blais and Bodet 2006). 'Neither theory demands that all voters have identical and perfectly accurate information (…) It is sufficient to assume that the issue stands actually taken by the parties have some impact on the images of those stands held by voters' (Westholm 1997, p. 870). The use of individually perceived left–right orderings may thereby be a sufficient alternative to pull directional and proximity models apart.

Third, one potential solution to the lack on a one-dimensional common space might lie in the use of a multidimensional space (cf. Carroll and Kubo 2021) or of separate policy dimensions (e.g. Alvarez and Nagler 2004). To some extent voters are better equipped to meet the demand of collective transitivity on these separate policy dimensions: they are more likely to agree on the substantive meaning of these dimensions. Nevertheless, the existence of common policy scales should also be tested rather than assumed.

Fourth, our analyses suggest that the unidimensional common space is more likely to hold some conditions: (i) In party systems with relatively few parties, (ii) in moderately polarised party systems, and (iii) for relevant and established parties rather than small and challenger parties. The unidimensional common space might be a useful assumption under the most favourable conditions and for a subset of the party system. Yet, even under such favourable conditions, the common space should not be considered self-evident.

We cannot continue to assume that the left–right dimension is a 'common yardstick' (Van der Eijk and Niemöller 1983) that facilitates 'general communication exchanges' (Sani 1974), or that voters 'can reasonably recognise the parties' positions' (Jou and Dalton 2017). This interpretation of left and right is wrong. The problem is not only that voters lack a common understanding of party positions on the left–right scale (Bauer et al. 2017; Lindqvist and Dornschneider-Elkink 2023), but the incorrect assumption among scholars that voters do.

Appendix 1

Parties and elections included in the cross-sectional analysis (abbreviations according to CSES code book).

	Year of election	Main analysis (All parties)		Robustness check (> 10%)	
		Parties	*N*	Parties	*N*
Australia	2013	LP/LNP, ALP, AG, NP	4	LP/LNP, ALP	2
Austria	2013	SPO, ÖVP, FPO, Grüne, BZO, Neos, Team Stronach	7	SPO, ÖVP, FPO, Grüne	4
Brazil	2014	PT, PSDB, PMDB, PSB, PSD, PR, DEM	7	PT, PSDB, PMDB	3
Bulgaria	2014	GERB, KzB/DL, DPS, RB, NFSB/IMRO, BBZ, Ataka, ABV	8	GERB, KzB/DL, DPS	3
Canada	2011	Con, NDP, Lib, GP	4	Con, NDP, Lib	3
Czech Republic	2013	CSSD, ANO 2011, KSCM, TOP 09, ODS, Usvit, KDU-CSL	7	CSSD, ANO 2011, KSCM, TOP 09	4

	Year of election	Main analysis (All parties)		Robustness check (> 10%)	
		Parties	N	Parties	N
Finland	2015	KESK, KOK, PS, SDP, VIHR, VAS, RKP, KD	8	KESK, KOK, PS, SDP	4
France	2012	PS, UMP, FN, FDG, MoDem, EELV	6	PS, UMP, FN, FDG	4
Germany	2013	CDU, SPD, Linke, Grüne, FDP	5	CDU, SPD	2
Great Britain	2015	Con, Lab, UKIP, LD, GP	5	Con, Lab, UKIP	3
Greece	2012	ND, SYRIZA, PASOK, ANEL, LS-XA, DIMAR, KKE	7	ND, SYRIZA, PASOK, ANEL	4
Hong Kong	2012	DAB, CPP, DP, PP, FTU, LAB, NPP, LP	8	DAB, CPP, DP	3
Iceland	2013	Sj, F, Sam, Graen, BF, Pi	6	Sj, F, Sam, Graen	4
Ireland	2011	FG, Lab, FF, SF, GP	5	FG, Lab, FF	3
Israel	2013	L/YB, YA, MHH, HH, Shas, Hat	6	L/YB, YA, MHH	3
Japan	2013	LDP, DPJ, JCP, YP, JPR, NK	6	LDP, DPJ, JPR, NK	4
Kenya	2013	TNA, ODM, URP, WDM-K, UDFP, NARC-K	6	TNA, ODM, URP	3
Mexico	2015	PRI, PAN, PRD, MORENA, PVEM, MC, PT	7	PRI, PAN, PRD	3
Montenegro	2012	CG, DF, SNP, PCG, BS, PZJ-FPB, AK, HGI	8	CG, DF, SNP	3
The Netherlands	2017	VVD, PvdA, PVV, SP, CDA, D66, GL, CU, SGP, PvdD	10	VVD, PvdA, PVV	3
New Zealand	2014	NP, Lab, GP, NZF, MANA, MP, ACT, UFNZ	8	NP, Lab, GP	3
Norway	2013	AP, H, FRP, KFR, SP, V, SV, MDG	8	AP, H, FRP	3
Philippines	2016	LP, NPC, NP, UNA, PDP-Laban	5	LP, NPC	2
Poland	2011	PO, PiS, RP, PSL, SLD	5	PO, PiS, RP	3
Portugal	2015	PaF, PS, BE, CDU, CDS-PP	5	PaF, PS, BE	3
Romania	2012	USL, ARD, PP-DD, UDMR	4	USL, ARD, PP-DD	3
Serbia	2012	SNS, DS, SPS, DSS, LDP, URS, SRS	7	SNS, DS, SPS	3
Slovakia	2016	Smer, SaS, OLaNO, SNS, LsNS, SR, MH, S	8	Smer, SaS, OLaNO	3
Slovenia	2011	LZJ-PS, SDS, SD, DIGV, DeSUS, SLS, N.Si, SNS	8	LZJ-PS, SDS, SD	3
South Africa	2014	ANC, DA, EFF, IFP, NFP, UDM, FF+, COPE, ACDP	9	ANC, DA	2
South Korea	2012	NFP, DUP, UPP, LFP	4	NFP, DUP, UPP	3
Sweden	2014	SAP, M, SD, MP, C, V, FP, KD	8	SAP, M, SD	3

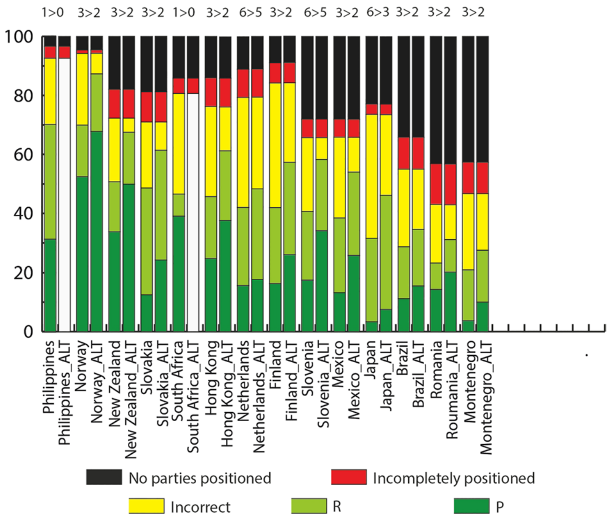
	Year of election	Main analysis (All parties)		Robustness check (> 10%)	
		Parties	*N*	Parties	*N*
Switzerland	2011	SVP, SP, FDP, CVP, GPS, GPL, BDP	7	SVP, SP, FDP, CVP	4
Thailand	2011	PPT, DP, BJT, Rak Thailand, CP, Chart Pattana, Rak Santi, Matubhum, Phalang Chon	9	PPT, DP	2
Turkey	2015	AKP, CHP, MHP, HDP	4	AKP, CHP, MHP, HDP	4
USA	2012	DEM, REP	2	DEM, REP	2

Appendix 2

Allowing errors between parties with highly similar averages.

The lack of support for *R* and *P* may be due to parties that in their average are positioned closely together. This issue is tackled via the Mokken scale analyses. Yet, we perform an additional analysis that departs from the idea that such differences perhaps should not factor into the test of *R* and *P*. Of course, that idea implies that the aggregation of voter perceptions of party positions cannot be reliably assumed for parties that are positioned closely together.

We ran additional analyses on the data set of relatively large parties (> 10%, Fig. 2). We assumed small differences in aggregated positions (≤ 0.5) to be irrelevant to the test of *R* and *P*. Under that assumption our findings suggest a few things (see Figure below).



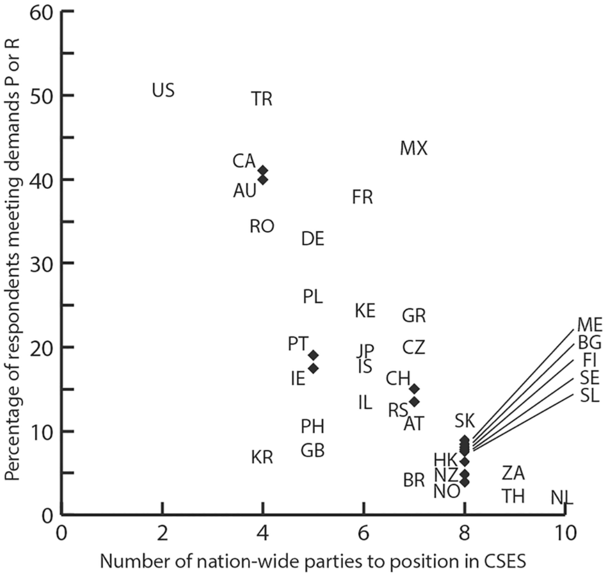
First, for the majority of the countries, party positions are so far apart that a check is not even possible. Second, in two of the fourteen countries where party distances are that small, there are already only two large parties; consequently, allowing their relative ordering to go in any direction makes any test of *P* and *R* simply impossible. The closeness of party positions among voters may even indicate a lack of familiarity with the scale, making any interpretation even more difficult. Third, in many countries this check makes only a relatively small difference. However, in some countries this check leads to substantially more support for the assumptions: Norway, New Zealand, Hong Kong, Finland, Slovenia, and Mexico.

The problem in deciding on the improvement of *R* and *P* in this way, is the lack of model fit indicators. When we release the number of demands (from 3 orderings to 2 or 1, or from 6 to 5 or 3), we can safely expect that demands *P* and *R* will automatically find more support. However, whether that support is sufficient to conclude that there may be a common space under this condition (i.e. large parties with large gaps between them), is unclear.

Mokken scaling allows a more direct test of ordering that takes the size of errors into account, albeit in a different way.

Appendix 3

Individual orderings of political parties by number of parties to position, CSES 2011–2016.



Declarations

Conflict of interest On behalf of all authors, the corresponding author states that there is no conflict of interest.

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