Europe's Common Left-Right Space

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Abstract

This study presents a new method to estimate the locations of voters, parties, and European political groups in the same ideological space using left-right placements by voters. We apply our method to the 2009 European Election Survey and demonstrate that the improvement in party estimates that one gains from fixing various survey bias issues is significant. Our scaling strategy provides left-right positions of voters and party positions for 162 parties — more than traditional expert survey studies currently provide. We test the convergent validity of these positions in multiple ways and demonstrate how rescaled voter and party positions can be used in cross-national research.

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1 Introduction

The study of elections and party competition is nowadays unthinkable without comparative measures of the ideological positions of voters and political parties, and a significant body of research attempts to quantify these positions along the principal left-right dimension of ideological conflict. One important source of data used in such estimates are voter surveys in Europe, which frequently include questions that ask voters to place themselves and various political parties on an abstract left-right scale. However, comparisons of voter and party positions estimated using survey data is complicated in systematic ways, notably by voter perceptual bias and issues relating to comparability across different countries. In this paper we propose an alternative that corrects for these issues, allowing voter and party positions to be placed in a common cross-national left-right space.

Using voter surveys to locate parties entails specific challenges. Notable limits to such surveys include the fact that one cannot produce estimates of party shifts over time or on specific policy dimensions. Such challenges can surely be better addressed by other techniques (i.e. expert surveys or manifesto analysis). But surveys also enjoy some significant advantages over existing techniques. First, they cost little in the sense that the questions required to produce our estimates have regularly been asked of respondents in cross-national surveys like the European Election Study. Thus, no incremental effort to gather additional data is necessary. Secondly, surveys tend to produce estimates for a larger number of political parties than is typical of current expert surveys or manifesto analyses. Finally, and most importantly, surveys are specifically tailored to locate voter positions as well, making the data especially well suited to studies examining interactions between parties and voters.

Despite this potential, the use of perceptual data to locate parties presents a number of unique problems that have not yet been resolved. One central issue is the problem of dealing with systematic respondent-level bias, a problem more commonly known in the literature as differential item functioning (Aldrich and McKelvey, 1977; Brady, 1985; Palfrey and Poole, 1987; King et al., 2003; Alvarez and Nagler, 2004). Stated differently, if a respondent believes that party A lies to the left of party B, there are multiple ways this idea can be expressed — on an 11-point left-right scale, parties A and B could be placed at 1 and 2, or at 4 and 8 respectively. Secondly, biases in scale perception may also manifest themselves at the cross-national level. More specifically, if respondents in France place party A at a 4 and respondents in Bulgaria also place party B at a 4 on the same 11-point scale, does this necessarily imply that party A and party B occupy the same ideological position on the left-right scale?

Our scaling approach deals with both issues. First, we correct for systematic perceptual biases of survey respondents *within* countries to place parties and voters on the same national scale. Subsequently, we rescale country-specific estimates into a *common* cross-national left-right space by exploiting the fact that national parties affiliate with political groups in the European Parliament. Estimates of uncertainty are then generated through the use of the non-parametric bootstrap (Efron and Tibshirani, 1994). Our scaling strategy provides party positions for 162 parties — more than traditional expert survey studies currently provide — together with their standard errors, and, for the first time, comparable voter placements and left-right positions for the European political groups.[1]

Our paper proceeds in four stages. First, we identify and discuss common problems that appear in cross-national studies of voters and parties. Next, we discuss the specifics of our model, which combines earlier work by Aldrich and McKelvey (1977) and Groseclose et al. (1999). In our results section, we validate our estimates in several ways. First, we examine party and voter locations in the United Kingdom and demonstrate that rescaled estimates not only appear to be substantively consistent with prior expectations, while estimates that fail to correct for perception bias are not, but also that rescaled estimates significantly improve the model fit in a spatial model of voting with valence. Second, we test the validity of party position estimates against those available from expert surveys and find that the two correlate highly. Third, we calculate a measure of party system compactness (Alvarez and Nagler, 2004) and show that using rescaled estimates changes the rank ordering of countries despite the robust nature of the measure employed. Finally,

[1] All estimates and replication code will be made available on our website.

we replicate a study of individual government defection at European Parliament elections (Hobolt et al., 2009) and show using rescaled estimates improves the model and yields results that are accurately in line with their theoretical expectations. We conclude with a discussion of directions for future research.

2 Placing Voters and Parties on a Scale: Common Problems

Scholars face two challenges in using surveys to jointly estimate voter and party ideology. First, comparisons of ideological scales across different countries is difficult because respondents typically only locate parties within their own countries. In the absence of respondents who rate parties across countries on a common scale, linking ideological estimates across countries requires some combination of additional assumptions or data. Secondly, ideological estimates within countries are complicated by the fact that respondents often use scales in different ways. We discuss these two issues in greater detail below.

2.1 Comparisons across Countries

Comparative studies of party systems, voting behavior, or policy-making would not be feasible without accurate estimates voter and party positions on ideological scales. Most research tries to capture the conflict between parties on a principal dimension of conflict, often expressed as the "left-right" dimension of politics. The most frequently-used sources for cross-national data on party positions are the *Comparative Manifestos Data* (Budge et al., 1987, 2001; Klingemann et al., 2006) and expert surveys on party positions (Laver and Hunt, 1992; Benoit and Laver, 2006; Steenbergen and Marks, 2007; Hooghe et al., 2010). In both instances, cross-national comparisons are made possible by assumption. In the case of manifesto data, a common coding scheme as well as a common definition of "left-right" is applied to manifesto sources from all countries. In the case of expert surveys, one needs to assume that experts from different countries interpret the response scales presented to them in a similar fashion.

Yet, there is no straightforward way on how to combine data on party position with data on voter positions on the same scale. The most common approach is to transform the party position scale to the scale from surveys and and merge it with voter placements (e.g. Benoit and Laver, 2005; Hobolt et al., 2009; Duch et al., 2010). When calculating ideological distances from these data, part of the variation in the data may be explained by measurement error arising from combining the data in this way. An obvious alternative is to focus exclusively on surveys to estimate party and voter locations, which is possible as long as voters can perceive ideological differences between parties. This is likely to be the case on those dimensions on which parties compete in elections, and the left-right scale offers the most obvious choice. Other scholars have therefore stayed exclusively with voter surveys to compare voters and parties and use voters' perceptions of party locations (e.g. Blais et al., 2001; Kedar, 2005). Yet, the issue of how voter *perceptions of party locations* can be transformed into actual *cross-nationally comparable party positions* remains an unresolved issue.

Rather than making the assumption that respondents from different countries interpret the response scales presented to them in an identical manner, our estimation leverages an additional source of data to facilitate inter-country comparisons. In the European context, our goal is to rescale position estimates of national parties and voters into a left-right space that is common to all of Europe. This problem is complicated by the fact that many scales need to be rescaled simultaneously, as each country will have its own set of party placements. We address this issue by exploiting the membership of national parties in their respective political groups inside the European Parliament. The key idea justifying the use of European political group membership to link nationally-estimated ideological spaces into a common European space is that national parties choose their party group affiliations largely on the basis of left-right ideological conflict — an assumption that enjoys substantial empirical support in the work of Hix et al. (2007) and McElroy and Benoit (2010).[2]

[2]This idea follows a trend in the past decade to pay closer attention to how ideological estimates can be compared across different political institutions and actor groups. In the US context, scholars have proposed solutions to bridge the legislature, the presidency, and courts (Bailey, 2007), media outlets and legislators (Groseclose and Milyo, 2005), and media outlets and justices (Ho and Quinn, 2008). Common to these contributions is the idea that additional sources of data, more commonly known as "bridging observations", can be used to rescale estimates from different institutional settings into a

2.2 Scale Perception Issues in Surveys

Surveys, however, come with their own limitations. A well-known adverse feature of ideological response scales is that such scales can be perceived differently by individual respondents, leading to interpersonal incomparability of the answers (Aldrich and McKelvey, 1977; Brady, 1985; Palfrey and Poole, 1987; King et al., 2003; Alvarez and Nagler, 2004). In the context of ideological scales, the problem manifests itself in two ways. In a seminal article, Aldrich and McKelvey (1977) argue that respondents anchor the scales according to their own interpretation of the endpoints and interpretations of the intervals on the response scales. This implies that survey respondents may, in fact, agree where various parties stand on a left-right dimension, but because each respondent shifts and stretches the response scale, the reported positions deviate. As a result, perceptions of parties will differ, but this variation in party positioning may be explained to some extent by scale perception issues.[3]

To illustrate how scale perception issues might affect voter placements, suppose there are two voters who are asked to place three British parties, Labour, Liberal Democrats, and Conservatives on an 11-point (0-10) left-right response scale. The first voter places Labour at 1, LibDems at 3, and Conservatives at 4. The second voter locates Labour at 0, LibDems at 5, and Conservatives at 10. Thus, both voters use the scale in a similar way and locate parties in an identical rank order. However, each voter perceives the scale with very different levels of "bias" and "stretch". The first voter sees little ideological distance between the three major parties and believes they all lie far to the left. In contrast,

common space. For example, Poole (1998) exploits the fact that Congressmen often serve as senators to bridge ideological estimates in the U.S. House and Senate. Similarly, Bailey (2007) uses executive statements expressing approval or disapproval of various Supreme Court decisions to bridge ideological estimates of justices and legislators into a common space, and Groseclose and Milyo (2005) construct estimates of media outlet ideology by exploiting the propensity of media outlets and legislators to cite various think tanks.

[3]The problem is not limited to ideological response scales only. For instance, King et al. (2003) show that survey respondents in non-democratic China report higher levels of political efficacy than respondents in democratic Mexico. This paradox is due to the fact that Chinese citizens report higher levels of influence in government because they have lower standards for what should count as a satisfying level in any given response category. This response-category *differential item functioning* can be addressed by supplemental survey questions that provide a common reference point question with the same response categories. The answers to these "anchoring vignettes" can be used to rescale survey responses across different institutional settings into a common scale.

the second voter sees an enormous amount of distance between the three parties without the leftward bias of the first voter. In other words, the two voters might perfectly agree on where the parties stand; yet, their interpretation of the response scale leads them to placements that differ. The same response bias may apply to their own placement on the scale (Aldrich and McKelvey, 1977; Alvarez and Nagler, 2004).

3 The Model

We present an estimation approach that addresses all of these issues. For comparative scholars interested in cross-national comparisons of parties, our common space estimates allow different parties to be compared across countries on a left-right scale. Moreover, our approach is simple in that it relies on the standard questions found in many comparative surveys today and does not require additional questions. Finally, for scholars interested in cross-institutional research, we show how the concept of bridging observations can be exported to surveys, tailored here to the European context.

To obtain cross-national party position estimates, we use voter self-placements and their placements of political parties from the 2009 European Election Survey, but the approach can easily be applied to other comparative surveys such as the *Comparative Study of Electoral Systems* or previous *European Election Studies*. Our model estimates the ideological locations of parties in two stages. In the first stage, we apply the model developed by Aldrich and McKelvey (1977) to obtain ideological estimates of national parties and voters. These estimates correct for individual scale perception differences (differential item functioning) and are comparable *within* countries. Using these estimates, we then apply a technique adopted from Groseclose, Levitt, and Snyder (1999) to rescale those estimates into a common European left-right space using European Parliament group memberships as bridging observations. This generates voter and party placements that are cross-nationally comparable. Additionally, we recover ideological estimates of the European political groups in the same ideological space as auxiliary estimates that prove useful in validating our estimates. Finally, uncertainty estimates of party positions are generated via a non-parametric bootstrap.

3.1 Correcting for Individual Left-Right Scale Perception Differences

In the first stage, we estimate party locations within each country using the left-right placement question in the 2009 European Election Survey (EES, 2010; Egmond et al., 2010), which asks respondents to place various national parties on a 0-10 scale.[4] Assume there are J parties in a country to be placed on the scale by N respondents.[5] These parties each occupy a (true) latent position θ_j $(1 \leq j \leq J)$. Each respondent i $(1 \leq i \leq N)$ has a latent perception of the j-th party, defined as the true position with error distributed following standard Gauss-Markov assumptions, that is, $\theta_{ij} = \theta_j + \epsilon_{ij}$.[6] What the survey records, however, is only the observed perception on the left-right scale of party j by respondent i, Y_{ij}. Aldrich and McKelvey allow for differential item functioning to be accounted for by assuming that each individual has separate perceptual bias and stretch parameters α_i and β_i. These parameters distort the reports of respondent i's placement of party j such that:

$$\theta_j + \epsilon_{ij} = \theta_{ij} = \alpha_i + \beta_i Y_{ij}$$

Under this assumed model of behavior, the Aldrich-McKevley procedure jointly estimates the individual bias and stretch vectors $\hat{\alpha}_i$ and $\hat{\beta}_i$ and the party locations $\hat{\theta}_j$ by minimizing the sum of squared residuals across all respondents and parties such that

$$\sum_{\forall i,j} \epsilon_{ij} = \sum_{\forall i,j} \hat{\alpha}_i + \hat{\beta}_i Y_{ij} - \hat{\theta}_j,$$

subject to the model identification constraint that the estimate party positions $\hat{\theta}$ have mean zero and unit sum of squared distances from the mean.

The Aldrich-McKelvey technique is applied separately to each country survey from the European Election Study and produces estimates of left-right party locations that are comparable *within* each country. It also produces estimates of each respondent's latent

[4]The Aldrich/McKelvey technique assumes continuous scales. Readers who wish to employ similar techniques on data of a clearly ordinal nature are advised to consult Quinn (2004) for ordinal alternatives.
[5]For reference purposes, Aldrich and McKelvey refer to 'stimuli' not parties.
[6]Principally, this means independently distributed errors that are normally distributed with mean 0.

location in the same ideological space, θ_i, by transforming self-reported placements, X_i, with their individually-estimated bias and stretch parameters, such that

$$\hat{\theta}_i = \hat{\alpha}_i + \hat{\beta}_i X_i$$

These jointly scaled scores of voters and parties can significantly improve our position estimates, an argument we pursue further when applying our scores to a simple spatial model of voting with valence.[7]

3.2 Correcting for Cross-Country Differences

The aim of the second stage is to make voter and party locations comparable *across* countries. This is not possible with the first procedure alone, because while each country will likely have a different mean ideological location and variance, these parameters are assumed to be identical across countries under Aldrich-McKelvey. Let α_k be a country-specific shift parameter. Now suppose there are two countries that separately have α values of 0 and 0.5, but identical stretch parameters β_k. This implies that the mean of the parties on the left-right scale in the second country lies 0.5 units to the right of the mean position in the first country, so failure to account for this shift (i.e. by assuming $\alpha = 0$ for both countries) will bias our estimates of all parties in the second country by 0.5 units. In practice this would mean that, say, the German party mean position is assumed to be the same as the French party mean position, while in reality the party system in France may be shifted towards the left compared to the party system in Germany.

Facilitating cross-national comparisons therefore requires that each country's set of party locations be rescaled into a common space. To do this, we exploit the political group affiliations of each party in the European Parliament following the 2009 elections as cross-country bridging observations. Using the previously estimated location of parties as data, $(\hat{\theta}_{jkm})$, we assume:

[7]Notably, Palfrey and Poole (1987) use Monte Carlo simulation to show that the Aldrich-McKelvey procedure recovers party locations well, even if errors are heteroskedastic over stimuli.

$$\hat{\theta}_{jkm} = \psi_k + \gamma_k \theta_m + \epsilon_{jkm} \quad \forall j, k, m$$

where θ_{jkm} is the position of party j ($1 \leq j \leq J_k$) in country k ($1 \leq k \leq K$) belonging to European political group m ($1 \leq m \leq M$) as recovered in the first-stage Aldrich-McKelvey procedure. These scores are assumed to be functions of country-specific shift and stretch parameters ψ_k and γ_k, and the latent position of their corresponding European political group θ_m. We further assume that the error term ϵ_{jkm} is distributed normally with mean zero and variance σ^2, which allows estimation of our key parameters of interest through maximization of the likelihood function:

$$L(\psi_k, \gamma_k, \theta_m | \hat{\theta}_{jkm}) = \prod_{j=1}^{J_k} \prod_{k=1}^{K} \prod_{m=1}^{M} \phi(\frac{\hat{\theta}_{jkm} - \psi_k - \gamma_k \theta_m}{\sigma})$$

Identification of the model requires the constraining of two parameters. We accomplish this by constraining $\alpha = 0$ and $\beta = 1$ for a specific country, thus effectively placing all parties into the ideological space of that country.[8] Estimation of the parameters of interest is similar to the procedure proposed by Groseclose et al. (1999).[9]

Following estimation of all parameters, common space party positions are calculated by transforming each first-stage score as follows:

[8] As in all scaling problems, identification is strictly relative, so the choice of country is completely arbitrary. We ran our model by initially rescaling positions into the Bulgarian party space, and we again Z-transform all the scores to have mean zero and unit variance. The final estimates are not affected by the choice of which country's ideological space is chosen for the initial rescaling.

[9] Note that there are a few important differences. For Groseclose et al., θ_{jkm} are not party scores, but legislator ideal points obtained from the Americans for Democratic Action (ADA). Secondly, Groseclose et al. calculate standard errors for their adjusted ADA scores by inverting the Hessian of the equation above. This may potentially understate the true uncertainty of the adjusted scores in two ways. First, ADA scores are treated as data that are measured without error, yet they are simply ideal points calculated using no more than 30 roll call votes each year. Secondly, the model specified assumes that the error term for an individual at any point in time is uncorrelated with past or future errors. While this assumption may be true, it is noteworthy that other dynamic scaling techniques (e.g. Martin and Quinn, 2002) explicitly make the opposite assumption of autocorrelated errors. By scaling across countries, we avoid the second issue entirely, and we address the first issue by estimating uncertainty via the non-parametric bootstrap (Efron and Tibshirani, 1994) in both stages of estimation. Finally, a crucial difference between the two applications lies in the interpretation of θ_m. For Groseclose et al., θ_m is an individual meta-parameter that captures the mean ideal point of the legislator over time in the common space and is largely a "nuisance" parameter. In our application, the estimates for θ_m instead represents the locations of the European political groups in the common ideological space, a substantively important set of estimates that cannot otherwise be obtained from the European election survey data.

$$\theta_{jk}^{T} = \frac{\hat{\theta}_{jkm} - \hat{\psi}_k}{\hat{\gamma}_k}$$

Each respondent's self-placement in the common space is calculated analogously:

$$\theta_{ik}^{T} = \frac{\hat{\theta}_{ik} - \hat{\psi}_k}{\hat{\gamma}_k}$$

3.3 Generating uncertainty estimates

Following Efron and Tibshirani (1994), we are able to generate standard errors for our estimates using a non-parametric bootstrap. Bootstrapping is done by resampling survey respondents from the European Election Survey with replacement and reestimating both the national-level party estimates and the cross-national rescaling on the resampled data. We repeat this process over 100 iterations. Note that this simulates the uncertainty present in the respondent sampling process, but assumes no uncertainty in our knowledge about the European group affiliation of each party. We therefore assume that parties have sorted themselves into an ideologically compatible European group — an assumption that generally appears to be reasonable in most cases when we inspect our estimates.[10]

4 Results

We now discuss the model fit and present key results from the estimation that demonstrate that the rescaling yields more accurate estimates of voter and party placements from the surveys on a left-right scale. In a first step, we examine the nationally rescaled party and voter placements and apply them to a spatial model of valence in the UK. We demonstrate that model fit significantly increases using the rescaled left-right scale. In a second step, we examine the estimated party positions. While using placement data from the European Election Survey allows us to estimate positions for many more parties than currently available in expert surveys, we show that for those parties that appear in both in our estimation as well as in expert surveys the ideological estimates have a high convergent

[10] An important exception to our assumption of reasonable sorting is Estonia, which we discuss in greater detail later in the paper.

validity. We furthermore demonstrate how the party position and voter position data can be combined to calculate a measure for party system polarization. Finally, we apply an individual-level model of government defection at European elections using the rescaled dataset. The use of rescaled scores in the model improves model fit and has a substantive impact on one of the key explanatory variables in this model, a result that is in line with the theoretical expectations.

INSERT TABLE 1 HERE

4.1 Estimation Summary

The data are 11-point left-right self-placements of voters and of different parties from the 2009 European Election Survey.[11] Table 1 examines the summary statistics of our country-level estimates, the first part of the rescaling procedure. One immediate item to note is that a substantial fraction of the country-level samples disappear due to missing data issues. Recall that respondents only remain in the sample if they place themselves and all other parties on the left-right scale. While the European Election Study surveyed 1,000 respondents in each country, as many as 716 respondents get dropped in cases such as Bulgaria. This problem is likely to be particularly acute in countries where respondents are asked to place parties that are difficult to locate, resulting in survey non-response. Nevertheless, samples in all countries are sufficiently large for the estimation.[12] Each country-level estimation also identifies a set of survey respondents with negative weights — that is, respondents who see parties in a "mirror image" space where parties on the left and right are reversed. Palfrey and Poole (1987) demonstrate that these are

[11] This estimation is done for all countries except Malta. We omit Malta because as a two party system, the two country-specific parameters are uniquely identified. Due to unresolved coding issues in the data release of the 2009 EES affecting several countries, results from Belgium, Denmark, Sweden, and Spain are dropped from the joint rescaling.

[12] Saiegh (2009), for example, is able to estimate party locations in Costa Rica using as few as 31 respondents. An alternative way to estimate party positions using perceptual data is Poole's basic space procedure (Poole, 1998), which can be thought of as a generalization of the Aldrich-McKelvey technique to matrices with missing data and multiple dimensions. This technique has the benefit of retaining many observations that are discarded, but does not permit the estimation of voter ideal points in the same space. The latter is a significant issue in light of our use of the spatial model of voting later in this paper. Furthermore, we compared estimates for each country using Aldrich-McKelvey and Poole's Basic Space separately and found no meaningful differences — scores for every country in the sample correlated at $r = 0.98$ or above with the exception of Romania, which correlated at $r = 0.81$.

largely individuals with very low levels of political information. Building on this idea, they hypothesized that one reasonable measure of the political information for each respondent is the correlation between the individual's perceived location of the parties and the scaled party locations. We constructed an information measure from respondents by applying a standard two-parameter item response model to a battery of seven political information questions in the European Election survey, and found that our survey measure correlated with the Aldrich-McKelvey derived measure at $r = 0.28$. The moderate magnitude of this relationship is largely consistent with that reported earlier by Palfrey and Poole.

Two other fit statistics provide additional guidance in interpreting our model results. The reduction in variance is a ratio of the overall variance of perceptions in scaled data, divided by the average variance in the unscaled data. Substantively, it captures the percentage of variance that is corrected when differential item functioning is accounted for. These reductions range from approximately 112% of the variance in the original data in the case of Romania to roughly 5% for Italy. The R^2 statistic measures the percentage of variance in the scaled positions that can be explained by the left-right dimension.

4.2 Example 1: Voters and Parties in the UK

The summary statistics just described suggest a good statistical fit for our countries in the sample, but reveal little about the substance of those estimates. Figure 1 explores this issue by examining the estimates for the United Kingdom. On the left panel we show the recovered party coordinates overlaid on top of a density plot of rescaled voter ideal points. We find that the three major national UK parties (Labour, the Liberal Democrats, and the Conservatives) are recovered in an order consistent with prior expectations. Three smaller parties (Plaid Cymru, the Scottish National Party, and the Greens) are all located between Labour and the Conservatives, close to the Liberal Democrats. To the right of the Conservatives are the UK Independence Party and the British National Party, also consistent with prior expectations. The procedure is therefore able to recover party locations with survey data that is highly consistent with those obtained via expert surveys (Benoit and Laver, 2006; Hooghe et al., 2010). Stated differently, following the language of Campbell and Fiske (1959), our scores exhibit a high degree "convergent validity" in the sense that they are highly correlated with expert surveys while purportedly measuring the same concept.

INSERT FIGURE 1 HERE

While our rescaled scores exhibit good convergent validity after correcting for differential item functioning, unscaled scores do not always share the same properties. We calculate unscaled party locations by simply taking the mean party placement of each party on the 11-point scale, and plot our scaled scores against the unscaled ones with a regression line on the right panel of Figure 1. These scores not only differ significantly, but yield a completely different configuration of parties. While Labour, the Liberal Democrats, and the Conservatives are still aligned from left to right, there is substantial movement among all other parties. Under the unscaled means, the three smaller leftist parties (the Greens, Scottish Nationals, and Plaid Cymru) are all located to the left rather than the right of Labour. Furthermore, both the UK Independence Party and the British National Party are located to the left of the Conservatives. Expert survey data suggests, however, that the British National Party is to the right of the Conservatives (Hooghe et al., 2010). Even more distressing is the high degree of confidence that is implied by these estimates — each line on the plot captures the 95% confidence interval of each estimate, so we can reject the possibility that the UK Independence Party and the British National party is to the right of the Conservatives.[13]

While party locations recovered under Aldrich-McKelvey (AM) exhibit high convergent validity with expert surveys, the rescaling technique has the additional benefit of rescaling each survey respondent into the same ideological space. This allows a wide variety of theories regarding the spatial model of voting to be empirically tested (Downs, 1957; Enelow and Hinich, 1984). In applications of such models, the ideological distance between the voter and the party is typically a key variable of interest, and this distance can only be measured if both the voter and the party's ideal point are measured on the

[13]Standard errors for scaled party locations are derived from a non-parametric bootstrap, which is described in greater detail in the next section. For unscaled means, standard errors were calculated analytically.

same scale. We first estimate one such model here for the UK – a simple spatial model of voting with valence in one dimension.

Let i denote an individual who is considering voting for party j $(1 \leq j \leq J)$. Individual i has ideal point x_i, while party j has ideal point x_j. Assuming quadratic utility, the deterministic spatial utility that voter i receives for choosing party j is $U_{ij}^S = -(x_i - x_j)^2$. However, we also assume that each party has a valence parameter v_j that captures the non-spatial component of utility that each voter. The parameter v_j can substantively be thought of as the value of the party brand that it carries in the electorate regardless of its positioning on the left-right scale or the relevance of omitted spatial components orthogonal to the left-right dimension. Following the random utility framework of McFadden (1973), we can then specify the full utility that voter i gets from voting for party j as the sum of the non-spatial, spatial, and stochastic utilities, or $U_{ij} = v_j - (x_i - x_j)^2 + \epsilon$. If we further make the assumption that ϵ is distributed as a Type 1 extreme value distribution, then following Dhyrmes (1978) the probability that voter i chooses party j among the J possible party choices is:

$$Pr(V_{ij} = 1) = \frac{U_{ij}}{\sum_{k=1}^{J} U_{ik}} = \frac{e^{v_j - (x_i - x_j)^2}}{\sum_{k=1}^{J} e^{v_k - (x_i - x_j)^2}}$$

This is a conditional (multinomial) logit model with alternative and individual-specific variables, with v_j as the parameters of interest to be estimated using x_i and x_j as data. The parameter v_j is only identified in relative terms, so we constrain $v_j = 0$ for the Labour party. To simplify our model, we only retain voters who voted for one of the top four parties in the 2009 European election (Labour, the Liberal Democrats, the Conservatives, and the UK Independence Party). We construct our objective function based on the choice model described above and present two versions of our valence estimates in Table 2. In the unscaled estimate, x_i is simply the self-reported left-right location of the voter, X_i, and x_j is the mean placement of the party on the left-right scale by all voters, $\frac{\sum Y_{ij}}{N}$. In the AM estimates we instead use the party and voter locations shown on the left panel of Figure 1.

INSERT TABLE 2 HERE

Our estimates of the spatial model using both sets of estimates shows that the model fits using AM-derived estimates is considerably better, as the maximized log-likelihood is almost 300 points higher for a relatively small sample of $N = 218$. These likelihoods are directly comparable because the two models use the exact same parameters (i.e. the difference in degrees of freedom between the two models is zero). Model fits using the two different data sets also imply substantively different results — Liberal Democrats and the UKIP are estimated to have relatively powerful party brands using unscaled data, while the Conservatives are not. In contrast, our rescaled estimates imply that every party brand is powerful relative to Labour in the sense that they draw more votes that their spatial location alone would dictate — on a likelihood ratio test with 3 degrees of freedom against a null model with no valence parameters, we reject the null of no valence differences at $\alpha = 0.01$. Labour's valence disadvantage in European Parliament elections relative to all other parties in UK is consistent with theories of comparative political behavior and the notion of "second-order" elections (e.g. Reif and Schmitt, 1980) — as the party in government, Labour is likely to lose vote shares in any "second-order" European election, a topic we return to in the cross-national application of our estimates. Finally, our estimates are consistent with popular portrayals of a 2009 Labour government that was deeply unpopular in the wake of the 2008 financial crisis and a resurgent Conservative and Liberal-Democrat opposition.

4.3 Example 2: Cross-National Party Positions and Party System Polarization

In this section we discuss the cross-national party location estimates, which are obtained after rescaling the national party scores estimated under Aldrich-McKelvey under the assumption that parties belonging to the same European political group are more likely to share similar political preferences on the left-right dimension. We begin with an examination of our estimates and check for obvious estimation patterns and outliers. Next, we discuss some properties of our estimates. We find that our rescaled estimates demonstrate convergent validity with expert surveys. We then discuss the estimates of our auxiliary shift and stretch parameters, arguing that these contribute significantly to the fit of some countries and that they are consistent with prior substantive research findings. Next, we combine party positions with voter placements to calculate and compare a measure for party system polarization.

INSERT FIGURE 2 HERE

Figure 2 plots the distribution of party positions by European political groups using the rescaled estimates (top) and the simple mean positions from the survey (bottom). Once country-specific shift and stretch effects have been removed from the variation in party positions, the European groups look more compact than they do using simple means. Another way to compare the unscaled party mean positions with the rescaled estimates is to look at those parties that cross group lines. Specifically, we look at the the two major groups in the EP, the Group of the Progressive Alliance of Socialists and Democrats (S&D) and the Group of the European People's Party - Christian Democrats (EPP). Using unscaled positions, there are a number of S&D parties with a position more to the right than the most leftist EPP party.[14] The same is true for parties from the EPP that are to the left of the right-most party from the S&D.[15] However, this does not occur when we examine positions that have been rescaled using our two-step procedure. For these estimates, there are no cross-overs of parties from the two major EP groups.

Figure 3 presents the cross-national party position estimates separately for each political group and well as for parties that did not gain seats in the EP election or were unaffiliated with a group (not affiliated).[16] In general, the level of consistency between the left-right orderings of the national parties and their European group affiliation is very high — national parties that are more left-leaning than their rivals tend to affiliate with more left-leaning European groups. Variation in ideological heterogeneity on the left-right scale across European political groups is another important feature than appears

[14] These parties are: SPÖ (Austria), DP and MSD (Cyprus), SPD (Germany), SE (Estonia), Pasok (Greece), SDP (Finland), PS (France), Labour (Ireland), LSAP (Luxembourg), PvDA (Netherlands), PS (Portugal), Labour (UK).

[15] These parties are: KDU-CSL (Czech Republic), Unione di Centro (Italy), PSL (Poland), and UDMR (Romania).

[16] Scores have been Z-transformed after rescaling to allow for easier interpretation.

in our estimates. Notably, national parties belonging to the three left-leaning European political groups (EUL-NGL, Greens-EFA, and S&D) are much more tightly aligned (i.e. lower variance around European political group mean) than the right-leaning European political groups. This is largely expected for a Euroskeptic group like the EFD, but is more surprising for groups like ALDE.[17] Finally, parties that are not aligned with an EP group or parties that did not win any seats in the EP elections in 2009 are displayed under the category "no affiliation". As expected, these parties span the entire space, as they include parties from the far-left, center, and far-right of the political spectrum across Europe.

INSERT FIGURE 3 HERE

As a simple validation of our estimates, we compare our cross-national party estimates to those from the 2006 Chapel Hill expert survey (Figure 4). Our estimates correlate with scores derived from expert surveys at $r = 0.893$, suggesting a very high level of consistency. This convergent validity bodes well both for the expert survey literature as well as our estimates. Note, however, that relying on voter surveys can provide researchers with *more* party position estimates than expert surveys, as it the case in our estimation. We emphasize here that our estimates complement rather than replace expert surveys. Our technique will tend to perform well in cases where researchers wish to conduct research on a larger set of parties that are included as part of cross-national surveys but excluded from expert surveys. Our earlier valence example also suggests that our technique will perform well in situations where researchers wish to incorporate voter distances from parties as a variable in their analysis. Notably, our technique does not generalize well to obtaining estimates of party locations on specific issue dimensions, nor does it allow us to estimate party position changes over time.

INSERT FIGURE 4 HERE

After running a non-parametric bootstrap, we find that the mean standard error of our party estimates is 0.1. Since our rescaled estimates are Z-transformed, this implies

[17]But note the outlying ALDE parties, which are discussed in the paper.

that our standard error spans about 0.1 standard deviations of the European ideological space.[18] The magnitude of this standard error is slightly larger but in line with standard errors for ideal points derived by other scaling procedures such as Poole and Rosenthal's DW-NOMINATE (Lewis and Poole, 2004; Carroll et al., 2009).

One particular set of estimates in our data appears very unusual and requires further explanation. Estonia has two political parties that are members of the ALDE group (Eesti Keskerakond and Eesti Reformierakond), but these two parties occupy opposite ends of the political spectrum with other parties located in between them. Furthermore, it has a right wing party (Res Publica) that is a member of the EPP coalition, but lies to the left of the far right party (Eesti Reformierakond) despite being a member of a more right-wing coalition. This alignment is highly unusual, and our estimates suggest that a realignment of the Estonian parties or a change in membership in a European political group in the future is likely.[19]

As part of the rescaling process, we also obtain estimates of the locations of the European political groups themselves. An important point to note is that these estimates are obtained solely as by-products of the cross-national scales — unlike the national party locations which are in part obtained from voter placements of the parties, no voter placements of the European political groups were used to obtain these estimates.[20] These estimates are, of course, substantively important to European party research, but they also serve a useful purpose in checking the validity of our estimates. More specifically, if our estimation procedure is flawed it will not correctly recover the left-right configuration of the European political groups. We compared our estimates to the left-right placement of the groups obtained through expert surveys, published in Benoit and McElroy (2007).

[18]The stretch of our scale is of course determined by which parties are included in the European Election Survey. Figure 1 shows that this space includes no less than the five major parties in each country, but if one includes small extremist parties in various countries the range of the scale would likely be considerably larger.

[19]We conducted an additional test to determine if our estimator was in any way driving the unusual result in Estonia, plotting our recovered party locations against the mean placement of each party across all respondents. The rank ordering of the parties was unchanged after estimation and correlated with unscaled means at $r = 0.97$, but our estimator pushed the location of Eesti Keskerakond further to the left than the estimated location using unscaled means. This suggests some uncertainty about the actual location of Eesti Keskerakond, but it in no way undermines our claim that the alignment of parties in Estonia is highly irregular.

[20]In fact, such placements were not asked of respondents in the 2009 European Election Study.

These estimates are not ideal for comparison because they measure party positions during the 2004 European Parliament, whereas our estimates are drawn from 2009 European election survey. One important consequence of this is that two right-wing groups that existed in 2004 (the UEN and EDD) no longer exist in 2009, and hence cannot be compared.[21] However, using the five political groups that are directly comparable across elections, our 2009 estimates correlate with McElroy and Benoit's 2004 expert survey estimates at $r = 0.95$.[22]

INSERT TABLE 3 HERE

In addition to checking the validity and efficiency of our estimator, we are also interested in assessing the net effect our our rescaling. Stated differently, does rescaling actually change our estimates of party locations in a meaningful way compared to the estimation of party locations using simple means? We answer this question in three ways. First, we examine the country-specific stretch and shift estimates. Second, we compare rescaled party estimates with simple means. Third, we calculate and compare a measure of party system polarization. Table 3 provides estimates of the country-level rescaling parameters, $\hat{\alpha}$ (shift) and $\hat{\beta}$ (stretch). Two important patterns appear in the data. First, in 8 of the 21 cases shown we reject the possibility that the shift parameter α is equal to 0 at the standard 0.05 level of significance. Similarly, we reject the possibility that the stretch parameter is equal to 0 in 16 of 21 cases. The key to note is that while some

[21] McElroy and Benoit locate both the UEN and EDD to the right of the EPP, and our estimates of the new ECR party place them on virtually the same position as the EPP.

[22] One obvious extension of our model would be an application to the European integration question on the European Election Survey to generate a second dimension. One concern here is that because parties to a large degree align with European party groups on a left-right dimension (McElroy and Benoit, 2010), the party group membership would not serve as good bridging observations for a common European space. Our intuition on this appears to be correct — in replicating this procedure with the European integration question, there is virtually no difference in locations for every European Party group in our data except the EFD and the EUL-NGL, which were to the extremes on the Euroskeptic and pro-European ends of the scale. Furthermore, our estimates are largely bimodal, with a large group of EFD members on the Euroskeptic mode and all other parties clustered in a larger pro-European mode. Therefore, rather than using party group membership, we tested an alternative set of bridging observations: roll call votes of MEPs on constitutional issues (e.g. treaty reform). The problems here are the definition of a national party position (majority, two-thirds, unanimity?) and missing observations (if MEPs abstain on particular votes). In the end, using roll call votes, which were for the most part heavily lopsided on EU constitutional issues, we were not able to identify more than two "blocs" of a pro- and an anti-European camp of parties. In short, while the technique appears to identify which parties lie at which extremes of the European integration scale, the metric information that can be recovered through joint scaling appears questionable.

countries have a similar ideological distribution of parties, many do not. The substantive significance of the changes shown is quite large. Our estimate of the shift for Latvia for example implies that its mean party position is a full standard deviation away from that of Bulgaria, while our estimate of the stretch for the United Kingdom suggests that its parties span only 1/3 of Bulgaria's ideological range.

INSERT FIGURE 5 HERE

A second way to examine the impact of our rescaling procedure is by comparing the recovered coordinates to those obtained from the survey via simple means of left-right placements. In the context of our model, the simple means model not only implies no individual-specific bias and stretch parameters, but also a constant scale as well (i.e. a party rating of '5' in Bulgaria means the same thing as a party rating of '5' in the United Kingdom). Figure 5 visually represents this comparison with a simple scatterplot accompanied by a regression line of best fit. The two sets of estimates correlate well at $r = 0.827$, compared to the earlier reported correlation of $r = 0.893$ with expert scores. However, there are some extreme discrepancies between the two sets of estimates. Chief among these is our common space estimate of the British National Party. Our estimates rank the BNP the most right-wing of the 162 national parties in our data set. However, estimates using simple means suggests that the British National Party is much more mainstream, with 64 of the 162 parties lying to the right of the BNP. Substantively, we view this to be highly unlikely and take this to be further evidence in favor of our technique.

Our final way to check the substantive implication of our rescaling approach is done by aggregating information about voter and party dispersion. The comparative politics literature has a long tradition of examining the polarization of party systems on the basis of the ideological dispersion of parties (e.g. Taylor and Herman, 1971; Gross and Sigelman, 1984; Alvarez and Nagler, 2004; Sartori, 2005; Dalton, 2008; Rehm and Reilly, 2010). We choose to calculate one such measure offered in the literature by Alvarez and Nagler (2004). We choose this particular measure because it was developed to precisely take into account the scale perception issues in surveys discussed earlier. For Alvarez and

Nagler, the ideological differences between parties become comparable across countries in a measure of "the dispersion of parties in the issue space relative to the dispersion of voters in the same issue space" (Alvarez and Nagler, 2004, p.48). As a result, party system compactness (or polarization) is a function of three separate components. The first is the ideological dispersion of voters, the second is the ideological distance of the parties from a ideological center of gravity, and the third are the vote shares of the parties to take into account the relative size of the parties in the system. This measure for compactness of country k is calculated as follows (Alvarez and Nagler, 2004, p.50):

$$\text{COMPACTNESS}_k = \frac{\sigma_k}{\sum_{j=1}^{N} V_j |(P_{jk} - P_k)|},$$

where σ_k is the standard deviation of voter self-placements on left-right, V_j is the j-th party's share of the vote in the 2009 European elections, $P_j k$ is the placement of the j-th party on left-right, and P_k is the weighed mean of parties on left-right, where each party is weighted by its vote share. Alvarez and Nagler then argue that a large value of compactness "indicates that voters place *themselves* across a wide range of the issue space but the parties are clustered in a very narrow range of the issue space", suggesting a compact ideological space (Alvarez and Nagler, 2004, p.49). By incorporating both voter placements and party positions, this measure ought to be robust to scale perception issues. We examine this by calculating two versions. In the first version, we follow the original approach and input the original survey self-placements and the mean perceptions of the parties on left-right. In the second version, however, we use rescaled voter placements and rescaled party positions to calculate the measure.[23]

INSERT FIGURE 6 HERE

Figure 6 presents a scatterplot of party system compactness using the unscaled and the rescaled data. Because the measure is the ratio of voter dispersion to party dispersion, the measures are comparable and the line on the plot indicates if the two measures are

[23]We use the EES contextual dataset for the vote shares of the parties (EES, 2010; Czesnik et al., 2010). The total vote share covered in each country ranges between 64.04% in France and 99.99% in Austria and Luxembourg. The average total vote share of the parties is 88.64%. When using the rescaled data, we exclude respondents with negative A-M weights.

identical. The two sets of measures correlate highly at 0.71, suggesting a high robustness of this measure. Yet, the plot shows some important differences. For example, while the raw data suggest that Poland is the most compact party system relative to voters, this changes when using the rescaled data where Poland has the third most compact system after Romania and Slovakia. At the other end of the scale, the raw measure suggests that Czech Republic, Cyprus, and Hungary are the most polarized, whereas the rescaled data suggests that Austria, Cyprus, and France are. This means that while Alvarez and Nagler's party system compactness measure appears indeed robust in the majority of cases, rescaling the data does make a small yet potentially substantively significant difference in how party systems are ranked.

To sum up, we have provided repeated evidence that "low tech" measurement strategies such as simply taking the means of party placements have lower validity than our scaling strategy that maps voters and parties in a common ideological space. In the following section we demonstrate the extra leverage we gain from analyzing truly comparable party and voter placements on the left-right dimension.

4.4 Example 3: Government Defection in European Elections

Finally, we present an application that shows the advantage of using common scores for voters and parties for scholars of comparative political behavior. Existing measurement strategies that combine survey data with party position data from expert surveys or manifestos require strong assumptions for within and cross-country comparisons that are rarely questioned. Yet, it is quite likely that individual scale-perception biases and differential item functioning undermines the comparability and, therefore, the validity of those measurement strategies. Mapping voters and parties of different countries onto the same left-right dimension should facilitate empirical tests about the effect of the relative distances between parties and voters on political behavior. We expect two findings in this regard. First, first difference effects based on common scores should be stronger than for alternative approaches assuming comparability. Second, we expect to find an improved model fit, i.e., a higher log-likelihood when using our common scores. In the following, we employ our common scores to explain voting behavior in the 2009 European Parliament elections. Specifically, we apply an existing vote-choice model (Hobolt et al., 2009) to a fresh data set, the European Election Study 2009.

Hobolt et al. (2009) test at the individual level the well-known argument that European elections are second-order national elections (Reif and Schmitt, 1980) that are determined by domestic factors such as voter distance on the left-right dimension and satisfaction with governmental economic performance. Additionally, they argue that European issues do play a role in EP elections. Using previous European Election Studies (1999 and 2004) Hobolt et al. provide evidence that voters who voted for a governmental party at the preceding national election are more likely to defect from this party in the next EP elections the greater the distance from this party is on the issue of European integration. They demonstrate the same effect for the distance between voters and parties on left-right albeit their results are less robust on this dimension than on the European integration dimension. We attempt to replicate the baseline defection model of Hobolt et al. (2009) to see, first, to what degree their conclusions travel to 2009 and, second, what the consequences are of using common space positions of voters and parties instead of the raw scores derived from expert and voter surveys.

The authors run a hierarchical logit model predicting defection as a function of domestic factors (government approval, assessment of economy), individual-level controls (age, social class, strength of partisanship) as well as two policy distance variables (see Hobolt et al. (2009) for further details on how those variables are coded). According to their theory, respondents should consider the EU-dimension in addition to the left-right dimension in choosing which party to vote for. The distance variables are measured as absolute distances between a respondents self-placement and the position of the party (according to Benoit-Laver expert survey data) she voted for in the preceding national election on the left-right as well as the EU-dimension.

INSERT Table 4 HERE

In order to compare the estimation results, we restrict the sample to those respondents for which we were able to generate a distance on the left-right dimension using the Hobolt et al. measurement strategy and our common space rescaled scores.[24] Finally, we can only include those respondents who place themselves as well as all the respective national parties on the left-right dimension. To sum up, this leaves us with $N = 3453$ observations to estimate the Hobolt et al defection model with two different strategies to operationalize the left-right dimension using 2009 EES data. Table 4 provides an overview about the number of observations, the share of defectors as well as the parties in government across all countries in the estimation sample.

The first two columns of the Table 5 reproduces the published results of Hobolt et al. (2009). Given that the authors use Benoit-Laver expert survey data to place parties on both, left-right as well as a EU-dimension we first follow their strategy to construct both distance variables. Our replication results of the Hobolt et al defection model for 2009 are reported in the third column. Finally, in the forth column we report the estimation results when using our DIF-corrected and comparable left-right scores to generate the distance between voters and parties on this common European left-right dimension.

INSERT Table 5 HERE

Comparing the third column with the results of Hobolt et al. in 1999 and 2004 shows that the impact of domestic factors on the probability to defect from a governmental party by and large travels to 2009 as well. Low satisfaction with the national economy does increase an individuals's propensity to defect from the party they voted for in the previous national election.

[24]First, we exclude all observations which got assigned negative weights during the rescaling. Second, we excluded all data from France because there are no Benoit-Laver scores on the left-right dimensions available. We wonder, though, how Hobolt at al. could include France. Given the description of their coding strategy there should be (excluding France) merely 22 countries in their models and not 23 as they report in their table 1. Moreover, for the governing party in France, the UMP, there are no Benoit-Laver scores available. Third, as explained above we have not yet included data for our common space rescaling procedure from countries such as Sweden, Belgium, Denmark and Spain due to unresolved data cleaning issues in the EES data. Moreover we excluded all observations from Malta because of its two party system, that makes already the first rescaling step impossible (Note that the two country-specific parameters for Malta are uniquely identified). In sum, these criteria leave us with observations right now of 21 different countries. Fourth, similar to Hobolt et al. we consider a party as a governmental party even if it left the government just before the election (e.g., the Hungarian SzDSz left the government in April 2009) while we have to exclude governmental parties if they are not included in ESS (e.g., ADK of Cyprus). Finally, while trying to maximize the number of countries in our model, our results are robust to the exclusion of observations from countries such as Latvia, Luxembourg, and Ireland because of concurrent (national or local) elections. It could be argued, that concurrent elections provide incentives for voters that are not comparable with the situation in countries without concurrent elections.

The results of both distance variables for 2009, on the left-right as well as the EU dimension, are more similar to 2004 rather than the 1999 results. The size of the estimated point estimates of the left-right distance drop considerably while the estimated standard errors do not. In fact, the estimated coefficient for the impact of the left-right distance on defection is about twice as large in 2004 than it is in 2009 using the same coding strategy based on Benoit-Laver expert survey data. Thus, it does matter for predicting defection how far away a voter is from the party she voted for in the last national election. These results for the distance on left-right stay in stark contrasts to the 2004 results, which more clearly support the Hobolt et al. claims that in addition to domestic influences, vote choice in European Parliament elections are about Europe issues as well.

Moreover, Table 5 shows, as expected, an improved model fit when using our common scores in the fourth column rather than left-right distance measures based on Benoit-Laver expert data. While using the Hobolt et al strategy does slightly improve the fit by two points to -1571 when including such a left-right measure as compared to a baseline model without such a variable, the increase in model fit is three-times as large. The respective log-likelihood increases by 6 points as compared to the baseline model (with a log-likelihood of -1573 — not shown in the table).

Furthermore, we assess the consequence of using our estimated common space scores to generate absolute distances between voters and parties on the left-right dimension instead of the Hobolt et al measurement strategy. Given that for theses scores we no longer have to assume comparability within and across countries our measurement strategy should come with less measurement error. Comparing the results in column three and four that use the same data and are identical except for the operationalization of the distance on the left-right dimension shows that the size of the estimated coefficient is three times as large when using our common scores (column four) while the estimated standard errors are comparable across both models. Moreover, the model fit further improved as it can be seen when comparing the log-likelihoods across both columns. Substantively, when using our common scores for 2009 we find that the distance on the left-right dimension is positively related to an individual's probability to defect. Put differently, using rescaled scores the probability to defect increases by 5.9 percentage points (with a standard error of 1.8) when moving the distance variable from the 5^{th} to its 95^{th} percentile while fixing all other variables at their means, whereas this effect is only 4.9 percentage points (with a standard error of 2.2) for the unscaled data when moving again the distance variable from the 5^{th} to its 95^{th} percentile with all other covariates fixed at their means. To sum up, our replication exercise shows that assuming instead of estimating a common left-right dimension does come with a price tag that scholars should be aware of and take into account when designing studies that involve measurements that should be comparable within and particularly across different national contexts.

5 Discussion and Conclusion

In this paper, we introduced a new procedure designed to estimate voter and party locations across Europe in a common left-right space using a readily available data source. Relying solely on survey data, our technique produces estimates that can be compared across countries while correcting for various issues related to scale perception differences. Standard errors of our estimates can be generated easily via the non-parametric bootstrap. Our procedure has several advantages. In contrast to expert surveys our procedure provides party positions of a broader range of parties, specifically party positions for smaller parties that are typically excluded from ratings on expert surveys. All in all, we provide ideological party positions for 162 parties. Scholars of European politics benefit from our procedure because we provide comparable ideological positions for all European political groups within the same ideological space, and these ideological positions of the European political groups are solely a by-product of our estimation. Furthermore, we are able to validate our estimates in multiple ways. Our estimates correlate strongly with estimates obtained via expert surveys, and estimates of the European political groups exhibit similar levels of convergent validity. Moreover, the improvement in party estimates that one gains from fixing various scale perception issues is significant — in estimating a valence model for voters and parties in the United Kingdom, our corrected estimates provide a superior model fit to party estimates obtained from naive means of voter placements, and the same is true for a cross-national model of government defection in European elections.

Our technique to estimate party positions from surveys into a common space can be adapted in regions outside Europe as long as one is able to find appropriate "bridging observations" that help to glue together those underlying scales across countries. While differential item functioning correction via Aldrich-McKelvey rescaling is relatively straightforward for parties within the same country, for cross-national rescaling we leverage each party's affiliation to one European political group as a bridge in order to identify a common ideological space for the chosen context. While the European Union not only has a large number of party groups, it also has a fairly even distribution of membership across party groups within each country. We are convinced that some adaptation of our technique can produce similar cross-national estimates using other comparative surveys.

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Table 1: **Overall Fit of Left-Right Scales:**

Country	Respondents	Number Negative	Parties	Reduction in Variance	R^2
Austria	526	59	8	0.160	0.58
Bulgaria	284	31	8	0.173	0.56
Cyprus	749	40	6	0.092	0.71
Czech Republic	807	58	5	0.095	0.74
Estonia	453	87	6	0.232	0.57
Finland	849	48	8	0.117	0.63
France	611	28	8	0.063	0.73
Germany	875	50	5	0.097	0.74
Greece	764	51	6	0.127	0.67
Hungary	583	12	7	0.061	0.76
Ireland	738	132	6	0.699	0.39
Italy	605	27	8	0.048	0.78
Latvia	501	89	9	0.178	0.54
Lithuania	334	84	10	0.352	0.43
Luxembourg	601	27	8	0.216	0.53
Poland	367	37	6	0.340	0.5
Portugal	773	26	5	0.079	0.77
Romania	357	166	7	1.123	0.3
Slovakia	551	86	8	0.228	0.52
Slovenia	685	98	9	0.173	0.55
The Netherlands	695	108	11	0.136	0.55
United Kingdom	536	238	8	0.566	0.38

Note: Negative weighted respondents are those with low political information that perceive a reversed scale. Reduction in Variance measures improvement from scaled over unscaled scores, which can be interpreted as the amount of differential item functioning accounted for by the model.

Table 2: **Estimates of Valence Parameters in the UK, 2009 European Elections**

	Unscaled Estimate	AM Estimate
Liberal Democrat Valence	0.891 (0.262)	0.129 (0.207)
Conservative Party Valence	0.009 (0.421)	0.748 (0.19)
UKIP Valence	0.831 (0.304)	0.232 (0.220)
N	218	218
Log-Likelihood	-571.1114	-278.3912
Null Log-Likelihood	-580.598	-288.2430

Note: Valence for the Labour Party is omitted from estimation and fixed at 0. Estimates represent the non-spatial utility that each respondent gets for voting for that party instead of Labour, with standard errors in parenthesis. Unscaled estimates are calculated using mean party placements on left-right scale and unscaled respondent self-placements. The model shows a substantially better find using scores obtained via the Aldrich-McKelvey estimator. Null log-likelhoods are calculated from the same model with all valence parameters set to 0.

Table 3: **Estimates of country shift and stretch parameters ($\hat{\alpha}$ and $\hat{\beta}$) by country, 2009 European Elections:**

Country	Shift ($\hat{\alpha}$)	Stretch ($\hat{\beta}$)
Austria	0.06 (0.28)	0.37 (0.18)
Cyprus	0.91 (0.28)	0.53 (0.15)
Czech Republic	0.36 (0.27)	0.37 (0.13)
Germany	-0.32 (0.26)	0.41 (0.22)
Estonia	0.49 (0.20)	0.41 (0.16)
Greece	0.67 (0.27)	0.33 (0.12)
Finland	-0.81 (0.28)	0.47 (0.14)
France	-0.90 (0.28)	0.42 (0.13)
Hungary	0.00 (0.27)	0.61 (0.24)
Ireland	-0.92 (0.29)	0.51 (0.15)
Italy	-0.22 (0.22)	0.51 (0.18)
Lithuania	-0.06 (0.23)	0.73 (0.19)
Luxembourg	-0.32 (0.28)	0.91 (0.31)
Latvia	-1.03 (0.28)	0.38 (0.18)
The Netherlands	0.05 (0.27)	0.66 (0.22)
Poland	0.67 (0.26)	0.36 (0.11)
Portugal	-0.12 (0.25)	0.59 (0.21)
Romania	0.25 (0.22)	0.73 (0.21)
Slovenia	0.10 (0.28)	0.97 (0.25)
Slovakia	0.16 (0.21)	0.39 (0.12)
United Kingdom	0.17 (0.27)	0.33 (0.17)

Note: α and β are shift and stretch parameters facilitating comparison across European legislatures. Standard errors in parenthesis. Omitted reference category is Bulgaria, which is fixed to have $\alpha = 0$ and $\beta = 1$.

Table 4: **Defecting in the 2009 European Parliament Election**

Country	N	Defectors %	Parties in Government 2009
Austria	474	24.1	SP, VP
Bulgaria	130	30.8	DPS, NDSV
Cyprus	391	12.5	AKEL, DIKO, ADK*
Czech Republic	493	10.3	CSSD, ODS, SZ
Denmark	317	32.8	KF, V
Estonia	349	45.8	IRL, ERe, SDE-M
Finland	459	17.0	VIHR, KESK, RKP-SFP, KOK
France	248	13.7	UMP
Germany	516	21.1	CDU/CSU, SPD
Greece	336	22.3	ND
Hungary	252	17.5	MSZP, SzDSz**
Ireland	322	51.6	FF, Greens, PD
Italy	258	8.1	PDL, LN
Latvia	255	53.3	TB/LNNK, TP, ZZS, LPP/LC
Lithuania	207	9.7	LiCS, LRLS, TS-LKD
Luxembourg	353	25.8	CSV, LSAP
Malta	313	10.2	PN
The Netherlands	423	26.0	CDA, CU, PVDA
Poland	346	11.0	PSL, PO
Portugal	275	19.6	PS
Romania	442	13.6	PS-D, PD-L
Slovakia	406	8.1	SMER, SNS, HZDS
Slovenia	431	24.1	ZL-SD, LDS, ZARES, DeSUS
Spain	305	6.2	PSOE
Sweden	458	38.4	KD, M, FP, CP
United Kingdom	311	28.6	Labour

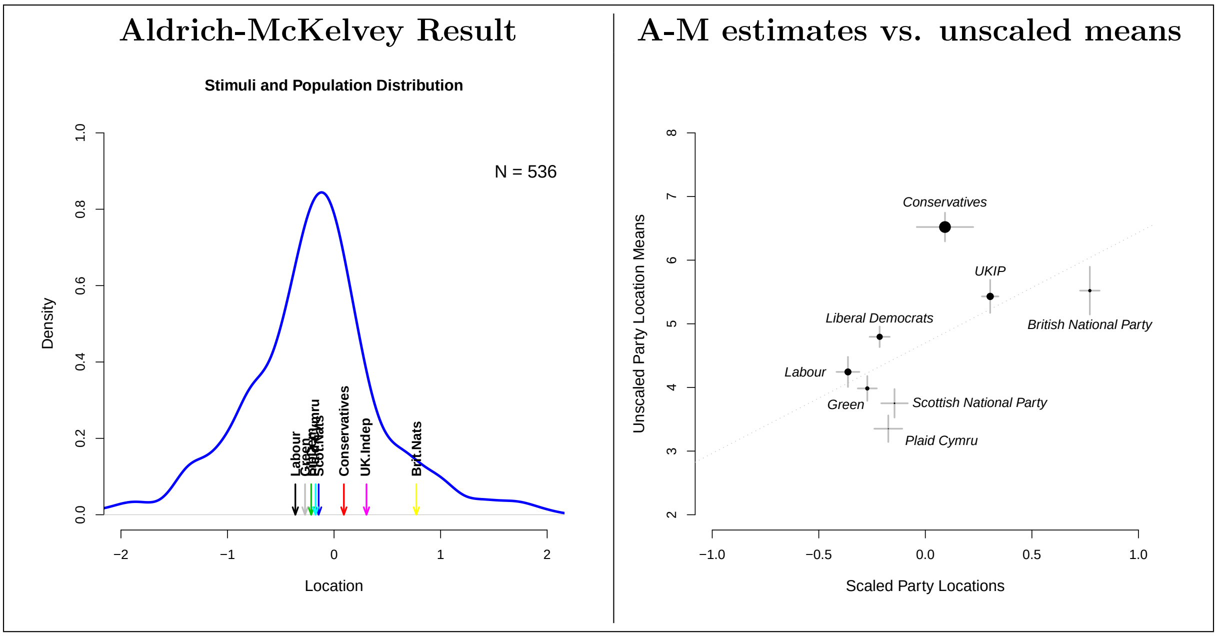
Source: 2009 European Election Study and ParlGov database (Dring and Manow 2010). All parties which held cabinet seats in June 2009 were treated as government parties.
* ADIK (CY) were not included in the EES survey.
** The Hungarian government was reshuffled in April 2009, as the SzDSz left the coalition, leaving the MSZP to form a minority government.

Table 5: **Explaining Government Defection in the 2009 European Elections**

	Hobolt et al		Benoit/Laver	Common Scores
	(1999)	(2004)	(2009)	(2009)
Age	-0.01**	-0.01**	0.01***	0.01***
	(0.00)	(0.00)	(0.00)	(0.00)
Social class	0.02	0.02	0.10**	0.11**
	(0.05)	(0.04)	(0.05)	(0.05)
Party identification	-0.57***	-0.63***	-0.85***	-0.85***
	(0.06)	(0.05)	(0.06)	(0.06)
Satisfaction with economy	-0.14**	-0.08	-0.13	-0.13
	(0.07)	(0.05)	(0.08)	(0.08)
Govt. approval	-0.34***	-0.95***	-0.27***	-0.26**
	(0.11)	(0.09)	(0.10)	(0.10)
Distance EU	0.04	0.05**	0.03	0.03
	(0.02)	(0.02)	(0.02)	(0.02)
Distance left-right	0.02	0.15***	0.08**	0.18***
	(0.03)	(0.03)	(0.03)	(0.05)
Intercept	0.18	0.01	-0.28	-0.34
	(0.37)	(0.31)	(0.30)	(0.30)
Log-Likelihood	-1399	-2147	-1571	-1567
No. of individuals	2868	4824	3453	3453
No. of countries	15	23 (?)	21	21

Note: Entires are hierarchical logit estimates with standard errors in parentheses. The first two columns reproduce results from Model 1 in Hobolt et al. (2009). Column three and four are based on EES 2009 data using either Benoit/Laver expert data or our common scores to operationalize party positions on left-right dimension. * $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

Figure 1: **Results of Aldrich-McKelvey Scaling for the United Kingdom, 2009.**



Note: Left panel show estimated locations of parties under AM rescaling, along with a density plot of estimated voter locations. Right panel plots estimated party locations recovered under Aldrich-McKelvey to scores obtained from taking the means of party placement scores with regression line. Bars represent 95% confidence intervals of each estimate, and size of points is proportional to vote share in 2009 European Parliament election.

Figure 2: **Distribution of Rescaled and Raw Party Positions in the EP, 2009**

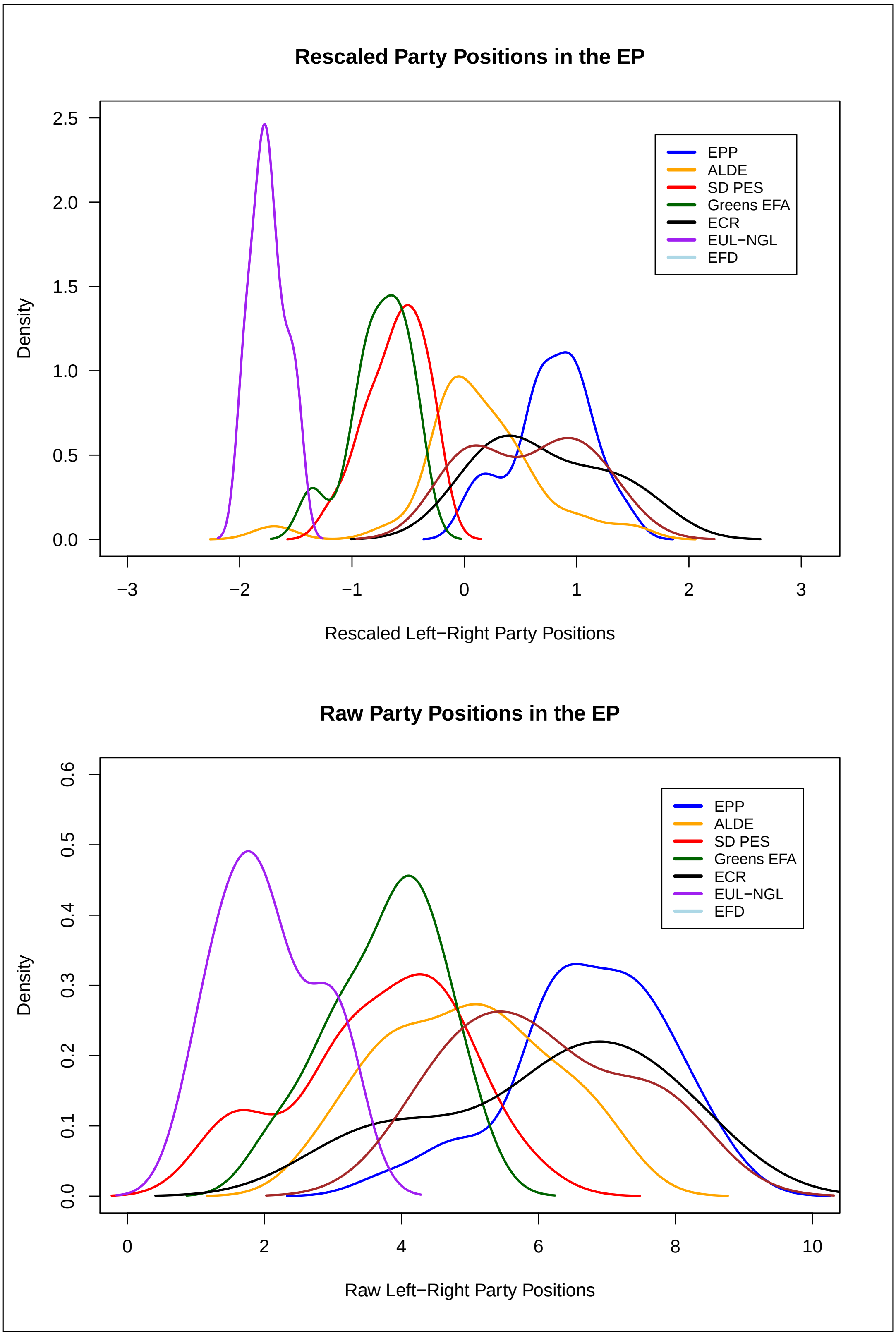


Figure 3: **European Parties in Common Space, 2009 European Elections.**

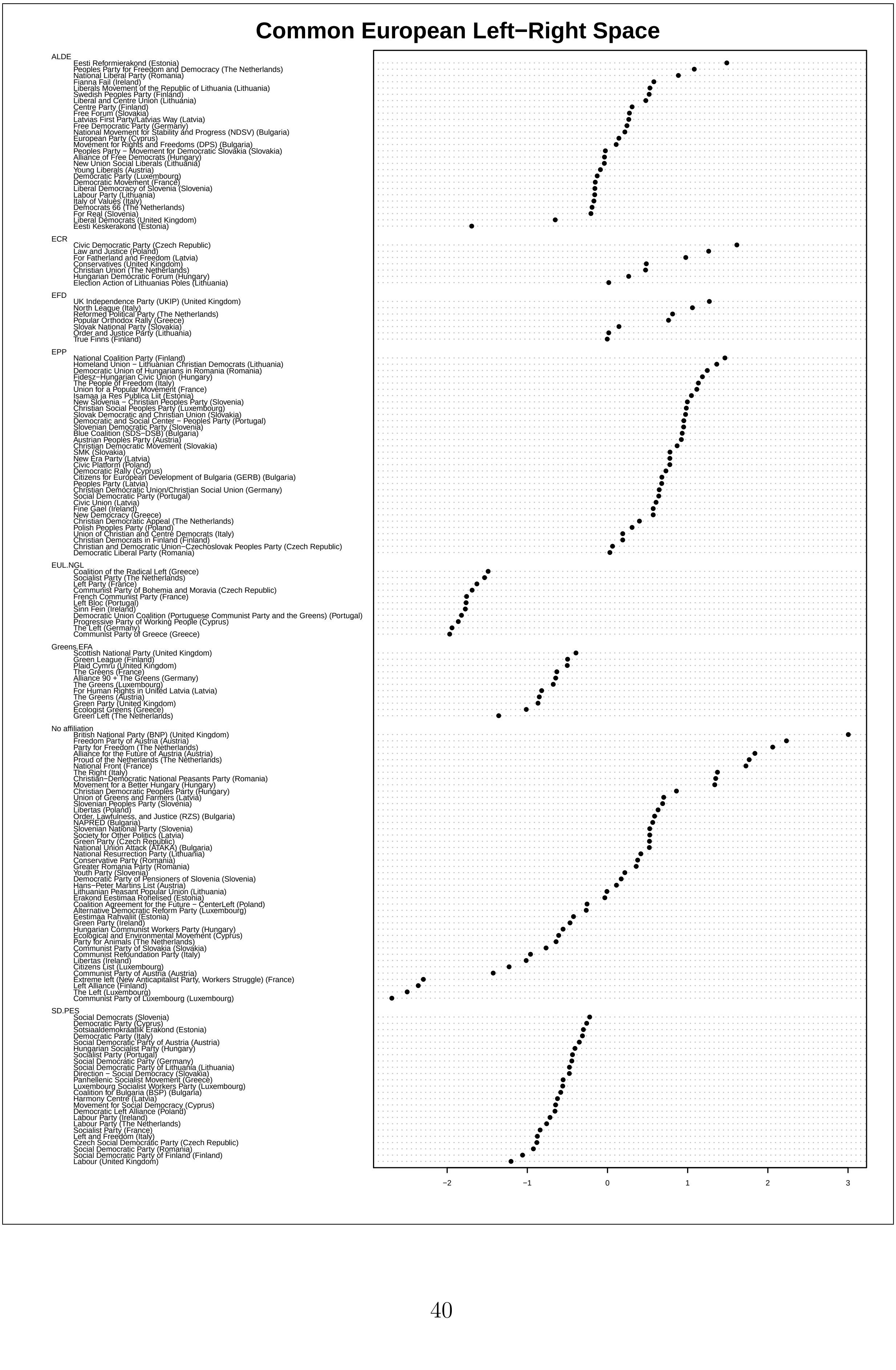


Figure 4: **Comparison of Rescaled and and Chapel Hill Expert Survey Placements, 2006**

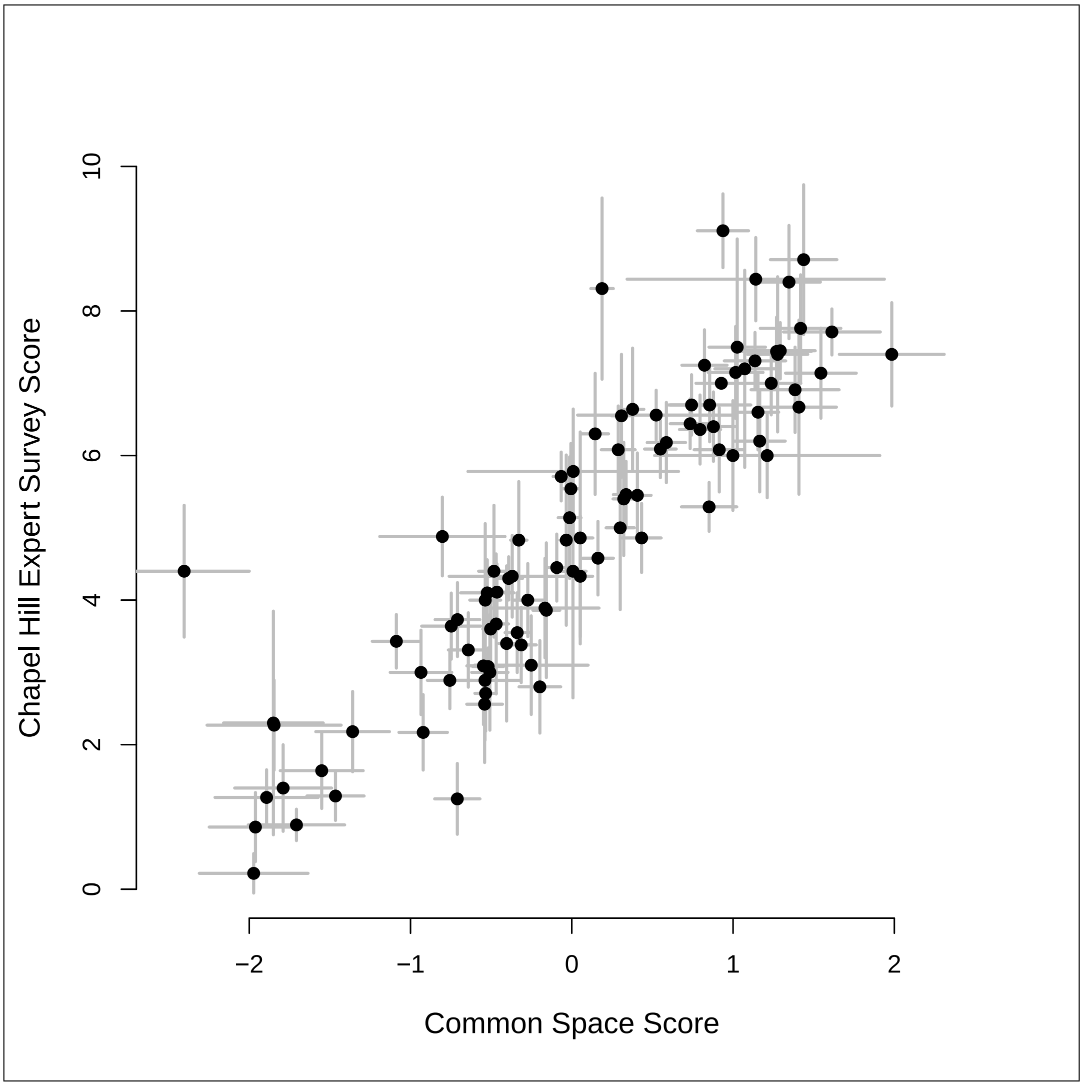
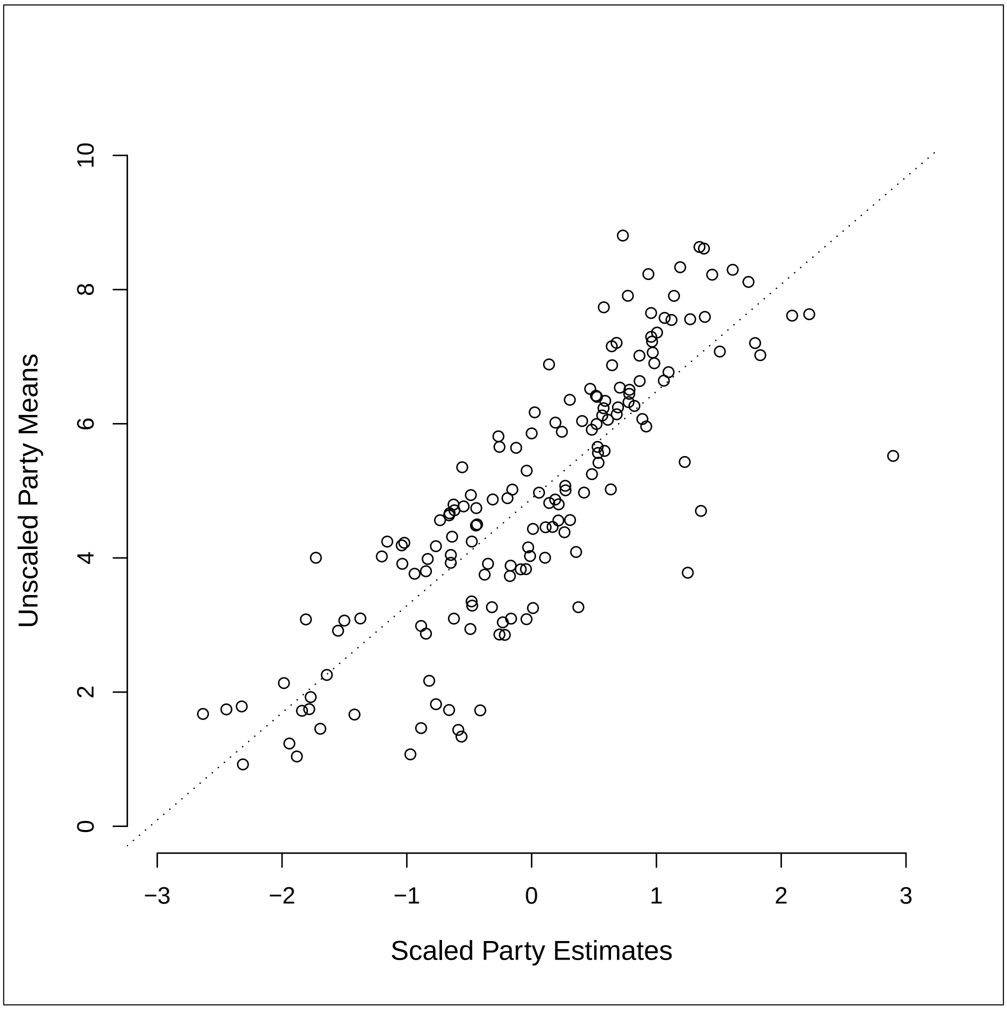


Figure 5: **Comparing Scaled vs. Unscaled Party Estimates, 2009 European Elections.**



Note: N=162 national party scores are shown in this comparison. The two estimates correlate at $r = 0.827$. Outlier to far right is the British National Party, which is the most right-wing party in Europe after rescaling, but ranks 98^{th} when placed using simple means.

Figure 6: **Party System Compactness Measures**