

Estimating Party Positions across Countries and Time—A Dynamic Latent Variable Model for Manifesto Data

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This article presents a new method for estimating positions of political parties across country- and time-specific contexts by introducing a latent variable model for manifesto data. We estimate latent positions and exploit bridge observations to make the scales comparable. We also incorporate expert survey data as prior information in the estimation process to avoid ex post facto interpretation of the latent space. To illustrate the empirical contribution of our method, we estimate the left-right positions of 388 parties competing in 238 elections across twenty-five countries and over sixty years. Compared to the puzzling volatility of existing estimates, we find that parties more modestly change their left-right positions over time. We also show that estimates without country- and time-specific bias parameters risk serious, systematic bias in about two-thirds of our data. This suggests that researchers should carefully consider the comparability of party positions across countries and/or time.

1 Introduction

Scholars use positions of political parties to analyze a variety of political phenomena, including the development of party systems, the (rational) behavior of voters, the spatial dimension of coalition formation, the composition of governments, the stability of cabinets, the spending for budgetary portfolios, the logic of legislative productivity and scrutiny, and last but not least the implications of bicameral settings (Warwick 1998; Tsebelis 1999; Martin and Stevenson 2001; Milner and Judkins 2004; Adams and Merrill 2006; Adams et al. 2006; Golder 2006; Tavits 2007; Ezrow 2008; Adams and Somer-Topcu 2009; Powell 2009; Maoz and Somer-Topcu 2010; Canes-Wrone and Park 2012). The common denominator of all of these studies is the investigation of political phenomena from a comparative perspective using positions of political parties that compete in distinct elections at a particular point in time (usually in national elections) within specific countries. Most often, scholars locate these parties applying a simplifying left-right scale, which indicates whether a political party favors more liberal or more state-oriented policies. For the empirical estimation of the positions of political parties, three data sources dominate the scholarly

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discussion: individual-level survey responses, expert judgments, and party manifestos (Hunt and Laver 1992; Lewis and King 1999; Budge et al. 2001; Glasgow 2001; Alvarez and Nagler 2004; Kedar 2005; Benoit and Laver 2006; McDonald, Mendes, and Kim 2007; Stoll 2010). Although individual-level citizen surveys and expert survey data locate political parties on predefined scales, scholars using manifesto data apply various methods to locate parties on scales *ex post facto*. These methods range from the computer-based analysis of word distributions to deterministic scaling techniques and latent variable models (Laver and Budge 1992; Gabel and Huber 2000; Laver, Benoit, and Garry 2003; Slapin and Proksch 2008; Lowe et al. 2011).

The current literature illuminates the advantages and disadvantages of each data source with relatively high cross-validity on the left-right dimension, but little is known about the comparability of the widely used scales, on which party positions are located across countries and over time. Conventionally, the implicit assumption is that the position of a political party on one scale from an election is comparable to a party's position in another election—independent from the fact that the elections take place at different points of time and/or in different countries. Because parties compete in specific elections, it is possible that the scales on which political parties are located differ by election. To account for potential biases emerging from country- and time-specific electoral contexts, we introduce a statistical model for estimating comparable party positions and combine expert survey data with manifesto data. Following the insights of the recent literature on the estimation of positions of legislators and political parties, we account for potential biases using additional model parameters and bridge observations (Groseclose, Levitt, and Snyder 1999; Bailey 2007; Shor and McCarty 2011; Gschwend, Lo, and Proksch 2012). One example where such biases can lead to misleading findings is the research that examines the effect of partisan ideology on public spending over time and across countries. A central hypothesis in this literature is that left parties in government increase spending, whereas right government parties cut it (e.g., Blais, Blake, and Dion 1993). Scholars that test this hypothesis using manifesto data without accounting for the potential biases risk drawing misleading conclusions about the role of ideology and government spending (Tavits and Letki 2009).

Compared to expert data, the comparative manifesto project provides data to estimate the positions of political parties over a long period of study and across many countries (Budge et al. 2001). The project codes the parties' electoral statements into policy categories, from which scholars infer positions of political parties using different methods (e.g., Gabel and Huber 2000; Budge et al. 2001; Bakker 2009; Lowe et al. 2011; Benoit et al. 2012). The theoretical expectation is that "parties compete by emphasizing policy areas they believe give them electoral advantage and by glossing over or ignoring those areas that they deem to help their rivals" (McDonald and Mendes 2001, 91). Since each manifesto is written for a distinct election, the question arises whether and to what extent the inferred positions of political parties can be compared across countries and over time without correcting for contextual electoral factors. For the widely used left-right scale, this means that country- and time-specific factors potentially distort the comparability of political parties' left-right positions. In other words, when a political party from a welfare state country and another party from a more (economically) liberal country mention the same policy category ten times for electoral purposes, it is questionable whether they defend the same position on a common scale. And when the same party mentions the same policy category ten times in the 1970s and 1990s, it is also questionable whether this party pursues the same position on a common scale.

The motivation of this article is to present a general and accessible method that pays attention to these potential country- and time-specific effects in the estimation of party positions from manifesto data. Our method builds upon Bayesian factor analytical models and incorporates a priori available information about the shape of the latent space. We also introduce two additional parameters to correct for biases that might arise from a priori incomparable data across countries and time. To account for the time dependency of party positions, we further model the dynamic movement of parties in the policy space with Legendre polynomials. When using this method to estimate party positions, scholars can be more confident about the comparability of the scales on which political parties are located. This could be a particular advantage for research that analyzes the interaction between actors and (horizontal and vertical) institutional settings by calculating spatial distances between parties across countries and/or over time.

We outline *one* empirical estimation strategy of our model that can be adapted to more specific research questions and theoretical models to be tested. In this application, we estimate the left-right positions of 388 political parties in twenty-five European Union (EU) member countries in the period 1945–2010. We incorporate expert survey data as prior information to preconstruct the left-right space. For a simultaneous estimation of country- and time-specific bias parameters, we identify bridge observations (Bailey 2007; Shor and McCarty 2011). We illustrate our estimates, which we call Manifesto Common Space Scores (MCSS), and evaluate their validity and robustness. Our scores offer three particular advantages over existing estimates: (1) MCSS are comparable across countries and time periods; (2) our scores are robust to changes in the empirical estimation strategy, which means that changes in the assumptions only modestly affect our findings; (3) our scores exhibit a high convergence validity with expert survey data and a high construct validity.

The remainder proceeds as follows. In the next section, we provide a brief overview of existing approaches to estimate party positions, which use manifestos as a data source. We then present our statistical model for coping with potential country- and time-specific effects. The fourth section outlines an application of the model, and the fifth section presents the results. In the sixth section, we assess the robustness and validity of our results. Finally, we discuss the implications of our contribution and provide a summary.

2 Existing Approaches for Estimating Party Positions

Expert surveys and party manifestos are the most prominent sources from which scholars estimate party positions. Although expert surveys use de facto knowledge of their interviewees on predefined scales, party positions are derived from party manifestos using two approaches: (1) methods relying on words as unit of analysis and (2) methods relying on coded manifesto data.

Two prominent methods exist for estimating the position of parties based on the distribution of words in a manifesto. Laver, Benoit, and Garry's (2003) Wordscore method computes the party positions using a known position of two reference texts (usually also manifestos). Their approach calculates the probability of reading one of the reference texts having seen a specific word in the analyzed text. Slapin and Proksch (2008) model the word count with a Poisson distribution and estimate the latent party position and various auxiliary parameters using maximum likelihood. Both methods are frequently applied to estimate left-right positions of political parties from party manifestos. Researchers also use these approaches to estimate the positions of other actors and/or dimensions by selecting other texts.

Several methods exist for estimating party positions from coded manifesto data. The ongoing Comparative Manifestos Project (CMP) provides these data by classifying and coding each quasi-sentence of a manifesto into a scheme of fifty-six categories, which capture a predefined set of political issues (Budge et al. 2001; Klingemann and Volkens 2007). Following the coding scheme of the CMP, the Euromanifestos project (EMP) (Braun et al. 2004) codes the content of party platforms for European Parliament (EP) elections in a similar manner. In the following, we focus on two distinct approaches to the extraction of parties' positions from these coded manifesto data: (1) deterministic scaling techniques and (2) latent variable models.

Deterministic scaling techniques are simple mathematical transformations of the quasi-sentence counts in the manifesto data. A prominent conventional transformation to a left-right scale is the so-called "Rile" scale, which subtracts the number of predefined leftists from rightist categories divided by the sum of these categories (Budge et al. 2001).¹ Because the coded manifesto data contain not only positional but also salience information, Lowe et al. (2011) propose a technique that is more suitable for estimating party positions on specific issues (e.g., environment or foreign affairs), which then can be aggregated to broader ideological dimensions (e.g., left-right). The authors specify a priori the poles of each policy issue or ideological dimension with at least two opposing manifesto categories. Each issue-specific position is then constructed by subtracting the logarithms of each pole's quasi-sentence count. Compared to previous scaling procedures

¹ In contrast to the Rile scale, Franzmann and Kaiser (2006) propose a procedure where the left-right dimension consists of different issues across countries.

assuming that the marginal effect of an additional quasi-sentence is constant (Laver and Budge 1992; Kim and Fording 2002), Lowe et al.'s (2011) technique can capture the decreasing marginal effect of quasi-sentences belonging to a category.

Latent variable models estimate statistically a party's position from the coded manifesto data rather than deterministically adding leftist and rightist positional categories. For placing political parties on the left-right dimension, Gabel and Huber (2000) are among the first who applied factor analysis to all fifty-six CMP categories. Their so-called vanilla method² inductively estimates the positions of political parties, which implies that the substantive meaning of the dimension has to be found *ex post facto*. One problem of this approach is that it leaves open the nature of the one-dimensional policy space to the *ex post facto* interpretation of the researcher. Benoit and Laver (2012, 216) thus recommend defining the political space *a priori* by exploiting common or expert knowledge. Bakker (2009) addresses this issue by proposing a Bayesian latent variable model with expert surveys' data as priors, in which the quasi-sentence count per category is modeled with a binomial distribution. More recently, Elff (2013) models the distribution of quasi-sentences across the CMP categories using the multinomial distribution. In contrast to Gabel and Huber (2000) and Bakker (2009), the author models the dependency between party positions with a vector autoregression function using maximum likelihood and reconstructing a party's position by empirical Bayes's techniques.

These different techniques and models for estimating party positions have stimulated a vivid methodological discussion about the advantages and disadvantages of each data source and method (e.g., Benoit and Laver 2007; Gemenis 2012; Grimmer and Stewart 2013). In general, the results of computer-based methods provide for high reliability but text selection by the individual researcher remains an open issue. Recently, Lowe and Benoit (2013) demonstrated how the results of methods modeling textual data can be validated on the basis of human judgments. Others are concerned with the intercoder reliability of the coded manifesto data (Mikhaylov, Laver, and Benoit 2012). The different estimates have been extensively cross-validated, but Dinas and Gemenis (2010) criticize that the plausibility of the estimates is hardly checked at the country level. Less disputed is the advantage of latent variable models, which reduce measurement error in comparison to deterministic scaling techniques.

To our knowledge, all authors—whether they use word counts or manifesto coded data with either deterministic or latent variable models—assume that their data on parties' positions are comparable across countries and over time. In our view, this is a strong assumption that only holds when the positions of political parties from different elections are located on the same scale. As party manifestos are written for a specific election, in which political parties position themselves relative to each other within a country at a particular point in time, we have good reasons to believe that a transformation is necessary for converting the election-specific scales into a common scale. We accordingly propose a Bayesian factor analytical model, which allows us to control for country- and time-specific effects.

3 The Statistical Model

Our model builds upon Bayesian factor analytical models (Quinn 2004; Jackman 2009), where the latent factors are the unobserved party positions on an ideological dimension. We adopt a Bayesian perspective to incorporate *a priori* available information about the shape of the latent space systematically as priors. As recommended by Benoit and Laver (2012), we thus avoid an *ad hoc*, *ex post facto* interpretation of the estimated positions. Generally speaking, there are three important differences between our model and existing latent variable models for coded manifesto data: (1) instead of assuming comparability of the manifesto data, we introduce two additional parameters to correct for biases that might arise from *a priori* incomparable data across countries and time; (2) we model the time dependency with Legendre polynomials rather than assuming

² In the following, we refer to the factor scores of the vanilla method as vanilla scores.

Table 1 Issue scales: each column lists the CMP/EMP categories (with their ID) used to construct the issue scales

<i>Issue</i>	<i>Pole A</i>	<i>Pole B</i>
Internationalism	109 Internationalism/negative	107 Internationalism/positive
European integration	110 European integration/negative	108 European integration/positive
National way of life	601 National way of life/positive	602 National way of life/negative
Military	105 Military/negative 106 Peace/positive	104 Military/positive
Freedom	201 Freedom and human rights/positive 202 Democracy/positive	605 Law and order/positive
Administration	404 Economic planning/positive	303 Governmental and administrative efficiency/positive
Enterprise	405 Corporatism/positive 412 Controlled economy/positive 413 Nationalization/positive	401 Free enterprise/positive
Market	403 Market regulation/positive	402 Incentives/positive
Protectionism	406 Protectionism/positive	407 Protectionism/negative
Macroeconomics	409 Keynesian demand management/positive	414 Economic orthodoxy/positive
Quality of life	416 Antigrowth economy/positive 501 Environmental protection/positive	410 Productivity/positive
Welfare state	503 Social justice/positive 504 Welfare state expansion/positive	505 Welfare state limitation/positive
Traditional morality	604 Traditional morality/negative	603 Traditional morality/positive
Multiculturalism	607 Multiculturalism/positive	608 Multiculturalism/negative
Labor groups	701 Labor groups/positive	702 Labor groups/negative
Target groups	705 Underprivileged minority groups/positive	704 Middle class and professional groups/positive

that the factors (the latent party positions) are independent; and (3) instead of using the original coded manifesto data we follow Lowe et al. (2011) and transform the coded data first.

We propose the following modeling strategy. We start with transforming the coded manifesto data (frequencies of quasi-sentences) into issue-specific positions, by following the transformation logic of Lowe et al. (2011).³ We take the data from König and Luig (2012) who construct sixteen positional issue scales from the coded manifesto data by assigning thirty-six of the fifty-six CMP categories to opposing poles of an issue-specific position.⁴ The sixteen issues scales are listed in Table 1 together with corresponding CMP categories. The supplementary appendix includes a more detailed description of the CMP categories used to construct the sixteen scales.

We continue with modeling the transformed data to extract the underlying party position. We define, \mathbf{y}_j to be the j ($j = 1, \dots, J$) row of our data matrix with L ($l = 1, \dots, L$) columns. Each cell contains the position of a party i ($i = 1, \dots, I$) that competes in the e^{th} election ($e = 1, \dots, E$) on a particular issue. Formally, our parameter of interest is $\phi_{(i,e,d)[j]}$ the unobserved D -dimensional

³ Although we apply the transformation of coded manifesto data recommended by Lowe et al. (2011), a more general modeling of the complete data-generating process would require modeling the assignment of categories to issue-specific scales, and then using them to extract the party's position on an ideological dimension. This would require an endogenous statistical selection between positional and saliency information.

⁴ Each issue-specific position is constructed by subtracting the logarithms of each pole's quasi-sentence count. As suggested by Lowe et al. (2011), we also add 0.5 to each pole, which "makes position estimates created from very small counts more stable, while barely affecting those derived from more reasonable numbers of sentences" (Lowe et al. 2011, 132).

Table 2 Parameters of the dynamic latent variable model

Name	Dimension	Description
$\phi_{i,e}$	$D \times 1$	Latent D -dimensional position of party i in election e
λ	$L \times D$	Matrix of loadings for L indicators (issue-specific policy positions)
ρ_c	$D \times 1$	Country bias parameters for country c
$\theta_{e,c}$	$D \times 1$	Time bias parameters for election e and country c
$\mathbf{b}_{k,i}$	$D \times 1$	k -th coefficient of the Legendre polynomials for party i
Σ	$L \times L$	Diagonal covariance matrix with elements w_l

($d = 1, \dots, D$) latent position of party i in election e . We here follow the Gelman-Hill notation (Gelman and Hill 2007) and let j select the corresponding indices i, e, k to the related the j^{th} observation, y_j , to a corresponding ϕ . Table 2 provides an overview about the parameters used in the model.

We assume that each indicator data point, $y_{l,j}$, can be decomposed into two parts: the latent position and the measurement error. The only difference across the indicators is the amount of measurement error. We suppose that the indicator variables are generated from a normal probability distribution of the form

$$y_j \sim N(\lambda(\phi_{(i,e)[j]} + \rho_{c[j]} + \theta_{(e,c)[j]}), \Sigma), \tag{1}$$

where $\phi_{i,e}$ is a vector of length D that contains the latent positions for i at e over D dimensions. ρ_c is a D -dimensional vector of country bias parameters unique to each country c ($c = 1, \dots, C$). Similar $\theta_{e,c}$ is a vector of length D with time bias parameters for each country c and election e . Σ denotes a diagonal covariance matrix with elements w_l^2 , of size $L \times L$ capturing the measurement error for each indicator. Finally, λ is the $L \times D$ matrix of loadings indicating how much a particular indicator is explained by the latent position. We follow Gabel and Huber (2000) as well as Bakker (2009) and assume that the factor loadings are constant across countries and time.⁵

The likelihood function with the standard normal density Φ is

$$\mathcal{L} \equiv p(\mathbf{Y}|\Gamma) \propto \prod_{l=1}^L \prod_{j=1}^J \Phi\left(\frac{y_{l,j} - \sum_{d=1}^D (\lambda_{l,d}(\phi_{d,(i,e)[j]} + \rho_{d,c[j]} + \theta_{d,(e,c)[j]}))}{w_l}\right), \tag{2}$$

with Γ being the collection of all parameters. The posterior density, $p(\Gamma|\mathbf{Y})$, can be stated as follows:

$$p(\Gamma|\mathbf{Y}) \propto p(\mathbf{Y}|\Gamma) \prod_{I,E} p(\phi_{i,e}) \prod_C p(\rho_c) \prod_{E,C} p(\theta_{e,c}) \prod_L p(\lambda_l|w_l^2)p(w_l^2). \tag{3}$$

Next, we specify the prior distributions, starting with $p(\phi_{i,e})$. Compared to previous approaches, which assume that the latent party positions are independent over time (Gabel and Huber 2000; Bakker 2009), we model the time dependency of party positions to use the entire indicator time series in our estimation. An advantage of modeling the dependence structure is that we can borrow strength from the complete time series and, hence, decrease uncertainty about the estimates. This approach reflects the common notion that parties' positions at past elections influence or constrain their future ones (e.g., Pierson 2000). We allow the latent positions to evolve over time by modeling them as polynomial functions of election instances. Similar to Poole and Rosenthal (1985), we capture the time dependency with Legendre polynomials. These Legendre polynomials are orthogonal in their codomain, which supports the estimation of latent positions.

⁵ It would be possible to allow for country and/or time-varying factor loadings, that is, to write equation (1) with $\lambda_{(e,c)[j]}$. However, we would need to make additional identifying assumptions to estimate such an extended model. Even if we only allow for time-varying loadings, we had to impose at least one equality constraint across the different time regimes. Since we are not aware of a theory guiding the selection of such an equality constraint, we leave to future research to explore this possibility.

We model each party position i in election e as follows:

$$\phi_{i,e} \sim \mathbf{N}\left(\mathbf{b}_{0,i} + \sum_{k=1}^K \mathbf{b}_{k,i} S_k(E_i, e), \mathbf{P}_{0i}\right), \quad (4)$$

where $S_k(E_i, e)$ is the k^{th} polynomial piece of the Legendre polynomials ($k = 1, \dots, K$) and \mathbf{P}_{0i} is a D -dimensional covariance matrix. E_i refers to the number of elections a party participated in. e refers to a specific election, and $\mathbf{b}_{k,i}$ to the D -dimensional coefficient vector to be estimated from the data. The larger the b coefficients in absolute terms, the more the party moves over time. We assume that the coefficients are drawn from a standard multivariate normal distribution of the form

$$\mathbf{b}_{k,i} \sim \mathbf{N}(0, 1). \quad (5)$$

In our estimation, we will use the first three Legendre ($K=3$) terms, which are as follows:

$$S_1(E_i, e) = (e - 1) - 1 \frac{2}{E_i - 1}, \quad (6a)$$

$$S_2(E_i, e) = \frac{1}{2}(3S_1(E_i, e)^2 - 1), \quad (7)$$

$$S_3(E_i, e) = \frac{1}{2}(5S_1(E_i, e)^3 - 3S_1(E_i, e)). \quad (8)$$

An alternative to the Legendre polynomials would be to assume a random walk structure for the latent positions (West and Harrison 1997; Martin and Quinn 2002). Although this option is very flexible, it comes at the cost that the position parameters are subject to large autocorrelation that increases the required MCMC run length. Another alternative would be to use simple polynomials such as linear and parabolic trends, that is, to use e, e^2, e^3 . We experimented with the different options and found that only the Legendre polynomials led to full convergence of the MCMC chain in a reasonable amount of time.

For the remaining parameters, we assume conjugate prior distributions as well:

$$\boldsymbol{\rho}_c \sim \mathbf{N}(0, \mathbf{R}_0), \quad (9)$$

$$\boldsymbol{\theta}_{e,c} \sim \mathbf{N}(0, \mathbf{T}_0), \quad (10)$$

$$\lambda_i | w_i^2 \sim \mathbf{N}(0, w_i^2 \mathbf{L}_0), \quad (11)$$

$$w_i^2 \sim \text{invGamma}(a_0, 1/b_0), \quad (12)$$

where the covariance matrices ($\mathbf{R}_0, \mathbf{T}_0, \mathbf{L}_0$) are assumed to be diagonal and of appropriate size.

4 Application

Our model for manifesto data is general and can be adapted to the specific needs of researchers. Using Monte Carlo experiments, we confirmed that our model is able to recover party positions unbiased from simulated data. In the following, we show how this model can be applied to extract latent positions of a large number of political parties from several countries on the left-right dimension, which is the most prominent, simplifying dimension used in many political science studies.

4.1 Data

We rely on the data of the CMP and the EMP considering twenty-five EU member countries in the period 1945–2010 (Wüst and Volkens 2003; Braun et al. 2004; Volkens et al. 2012).⁶ The estimation

⁶ We excluded Bulgaria and Romania (members of the EU since 2007) due to a lack of data about their national party delegations to the EP (see below). If two elections have been held in a country the same year, we dropped the data for the second election. That concerns Denmark (1953), the United Kingdom (1974), Ireland (1982), and Greece (1989).

Table 3 Descriptive statistics of the CMP/EMP data: Covered time period (col. 1), number of national parties and elections (cols. 2, 3), number of parties participating in EP elections (cols. 4, 5), and number of matches (col. 6) in the election used as a cross-country bridge (col. 7)

Country	Period	National elections		EP elections		No. of matches	Election
		No. of parties	No. of elections	No. of parties	No. of elections		
Austria	1949–2002	6	17	7	2	5	1995/1999
Belgium	1946–2003	22	19	17	6	4	1978/1979
Cyprus	1996–2001	7	2	5	1	4	2001/2004
Czech Republic	1990–2002	17	5	7	1	3	2002/2004
Denmark	1945–2007	18	24	11	6	6	1979/1979
Estonia	1992–2003	19	4	6	1	6	2003/2004
Finland	1945–2003	13	17	7	2	7	1995/1999
France	1946–2007	16	16	25	6	2	1978/1979
Germany	1949–2009	17	17	7	6	3	1976/1979
Greece	1974–2000	8	9	8	5	2	1981/1984
Hungary	1990–2002	11	4	4	1	3	2002/2004
Ireland	1948–2007	10	17	7	6	3	1977/1979
Italy	1946–2006	36	16	22	6	1	1979/1979
Latvia	1993–2002	24	4	7	1	6	2002/2004
Lithuania	1992–2000	15	3	10	1	5	2000/2004
Luxembourg	1945–1999	8	13	5	6	3	1979/1979
Malta	1996–1998	2	2	2	1	2	1998/2004
The Netherlands	1946–2003	14	18	14	6	3	1977/1979
Poland	1991–2007	32	6	8	1	6	2001/2004
Portugal	1975–2009	16	13	8	4	3	1987/1989
Slovakia	1990–2006	22	6	9	1	8	2002/2004
Slovenia	1990–2008	18	6	5	1	3	2004/2004
Spain	1977–2008	18	10	17	4	7	1986/1989
Sweden	1944–2010	9	21	8	2	7	1994/1999
United Kingdom	1945–2010	10	17	12	6	2	1979/1979
Σ		388	286	238	6	104	

of left-right positions is a challenging empirical task with our data set, which encompasses 388 national parties and 238 national elections. This task is even more demanding because our analysis also involves EMP data from the EU level. The EMP data have been collected by applying the CMP coding scheme to EP election manifestos ensuring the comparability of the data (Braun et al. 2004, 39). We accordingly consider 286 additional national party delegations competing for seats in six EP elections. This sample documents the richness of the manifesto data set for comparative research, which covers all elections after World War II until the most recent one in late 2000. Table 3 shows the number of elections and actors covered by our data set.

4.2 Identification

The parameters of our model are not identified, meaning that distinct parameter sets imply the same probability distribution for the data. In simple one-dimensional factor analytic models, one way to guarantee identification is to set two latent factors to different constants. The identification problem in our model is more serious, due to the presence of the two bias parameters for country and time. We attempt to solve this problem by considering three parameter sets (latent position, country bias, and time bias)—one at a time—while assuming we have knowledge about the other two. We show how each parameter set can be identified, whereby all three steps jointly identify the model.

We leverage two notions to identify our model: First, we exploit the functional equivalence of informed prior densities and parameter restrictions to place the party positions on a specific scale (Jackman 2009, 441). Second, we identify bridge observations to rule out invariance due to translation of the bias parameters. Bridge observations are known as instances when a party is known a priori to take the same position in two different elections (Shor and McCarty 2011). For our purposes, we assume that national parties defend the same position in their first EP election as in the previous national election. Furthermore, we maintain that parties with the highest relative seat share gains in their country take the same position in the next election. Below we discuss the motivation behind the two assumptions.

Similar to Bakker (2009), we use prior information to induce a scale for the party positions on the left-right dimension. Specifically, we define a scale for the party positions by setting each party's intercept and the variance ($b_{0,i}$ and P_{0i} in equation (4)) to the mean and variance of its party family. We calculate these means and variances for seven party families' attitudes on the left-right dimension using expert survey data by Hooghe et al. (2010) and Steenbergen and Marks (2007).⁷ We center all parties' priors that are not part of a party family at 0 and assign a variance of 11, which effectively provides no information about the location of the party in the empirical left-right space. This procedure only softly constrains parties belonging to the same party family to a specific segment of the left-right dimension, whereas the trajectory of this party on the segment is modeled with the Legendre polynomials. As Poole and Rosenthal (1985) do, we allow for up to three Legendre polynomials.

In contrast to our informed prior distributions on the latent party positions, we adopt vague prior distributions for λ , θ , and ρ : we center all three at zero and assign a variance of $R_0 = T_0 = L_0 = 10$. We also adopt a vague prior distribution for the γ distribution ($a_0 = 0.001$, $b_0 = 0.001$).

4.2.1 Cross-national bridges

A central question is whether and to what extent political parties are located on a common left-right dimension. To identify the country bias parameter, we exploit that many national parties also compete in transnational elections for seats in the EP since 1979. Assuming that the manifesto data from the same election are comparable, we use manifesto data on EP elections to estimate the size of the country bias. We justify our identification strategy with previous studies showing that in first-time EP elections, most elected candidates had a background in national politics (e.g., Corbett 1998, 71). Given this, we argue that parties took the same position in their first EP election as in the previous national election. We call this our "zero hour" hypothesis.⁸ This identification strategy allows to estimate the country bias, which is effectively the difference between the implied latent positions of the CMP and EMP data for the two elections. Table 1, column 7, lists this election set for each country and the number of parties assumed to take the same position in the elections (column 6).⁹ Fortunately for us, EMP data are collected and coded in a very similar manner to CMP. Hence, the EMP coding scheme consists of the same categories as the CMP, which allows us to calculate the sixteen positional issue scales for the EMP as we did for the CMP.

The identification of the country bias can be stated as follows: Let s be an indicator that takes the value 1 if the indicators (y_j) are coming from the CMP and 2 if they stem from EMP data. Denote the first EP election with e^\dagger and the previous national election with e^* . We then argue that

$$\phi_{i,e^\dagger} = \phi_{i,e^*}. \quad (13)$$

⁷ We consider the following party families: Left, Greens, Social Democrats, Liberals, Christian Democrats, Conservatives, and Nationalists. We use three survey waves: 1999, 2002, and 2006. To merge the manifesto data with the expert survey data, we use parts of the ParlGov data set provided by Döring and Manow (2012).

⁸ When Portugal and Spain joined the EU in 1987, and Sweden, Austria, Finland in 1995/1996, all five countries held special elections for the EP the same year. For our identification, we do not use these elections, but the regular EP elections in 1989 and 1999.

⁹ Not all parties that compete in national elections compete also in EP elections, and some only compete in EP election jointly in the form of party alliances. For our purposes, we only assumed that parties that compete individually in both types of elections take the same position.

To establish a reference point against which the ρ_c parameters in the CMP are determined, we set the country bias parameters for all EMP data to zero. Given knowledge about $\{\lambda, \phi, \theta\}$, ρ_c can now be uniquely determined. To see this, consider the following example for party i ¹⁰:

$$y_{s=1} = \lambda(\phi_{i, e^\dagger} + \rho_c + \theta_{e^\dagger, c}), \tag{14}$$

$$y_{s=2} = \lambda(\phi_{i, e^*} + 0 + \theta_{e^*, c}). \tag{15}$$

For all other EP elections, $e^{-\dagger}$, the positions of the national party delegations are modeled slightly differently, as in equation (4):

$$\phi_{i, e^{-\dagger}} \sim N\left(\phi_{i, e^*} + \sum_{k=1}^K b_{k, i} S_k(E_i, e), P_{0i}\right). \tag{16}$$

The intercept from equation (4) ($b_{0, i}$) is replaced with the position of the party in the previous national elections (ϕ_{i, e^*}). The intuition here is that the trajectory of positions from the national party delegations is “born” in the trajectory of national party positions and evolves from there independently.

An alternative could be to assume that the position of the national party is equal to the one of the national delegation to the EP when the national and European elections are closest to each other (“closeness” hypothesis). In these cases, the parties are likely to defend similar positions to remain credible. We will discuss the robustness of our estimates when using this assumption below. Another strategy might be to follow Gschwend, Lo, and Proksch (2012) and specify a function on how positions of the political groups in the EP are related to national party positions. These authors assume that national parties are linearly related to their respective political groups. However, this approach requires the estimation of additional parameters, which would further complicate the model setup.

4.2.2 Cross-election bridges

Similar to our “zero hour” hypothesis for country differences, we have to find bridge observations for coping with the potential time bias of each election. One assumption is that parties that gained the largest seat share in comparison to the previous election have little incentive to change their position in the next election (“incentive” hypothesis). To motivate this hypothesis, we refer to several scholars arguing that candidates are constrained in their ability to shift their policy position due to concerns about reputation and credibility (e.g., Bernhardt and Ingberman 1985; Ingberman 1989; Enelow and Munger 1993). Furthermore, Somer-Topcu (2009, 238) argues that parties that gained in an election become risk averse and will not shift their policy position “since any change would increase risks and uncertainty about the consequences of change.” On the other hand, parties losing vote shares at an election infer that the public opinion has shifted away from the party’s policies and, thus, have an incentive to adjust their position at the next election (Somer-Topcu 2009). Our argument is also in line with the Pavlovian win-stay, lose-shift strategy, which performs very well in simulations of the Prisoner’s Dilemma game (Nowak and Sigmund 1993). This identification strategy allows us to exploit the discrepancy between the implied latent position for such a party between two elections to estimate the time bias parameter.¹¹

More formally, let i^* be the party that gained the largest seat share in the e th election in a country c . We then assume that

$$\phi_{i^*, e} = \phi_{i^*, e+1}. \tag{17}$$

To establish a reference point against which all other time bias parameters, θ , are determined, we normalize $\theta_{1, c}$ to zero for all first elections in a country c . For any election e and the following

¹⁰ We omitted the stochastic component and the nested indexing for reasons of readability.

¹¹ For the first election in each country, when no party could gain seat shares relative to the previous election, we used the party with the highest seat share as a bridge observation.

election $e + 1$, there exist an i^* by definition, which allows us to identify all $\theta_{e,c}$ across all elections. To see this, consider as an example the equations for the party i^* that gained the largest seat share in the first election ($e = 1$) for some country c :

$$\mathbf{y}_e = \lambda(\phi_{i^*,1} + \rho_c + 0), \quad (18)$$

$$\mathbf{y}_{e+1} = \lambda(\phi_{i^*,1} + \rho_c + \theta_{2,c}). \quad (19)$$

Given knowledge about $\{\lambda, \phi, \rho_c\}$, $\theta_{2,c}$ can be uniquely determined.

The constraint requires again a modification of equation (4): (1) E_i needs to be discounted by the number of periods i won the largest seat share and (2) e now indexes only those elections in which i is not or was not i^* in the previous election. Let $t_{e,i}$ be this new index, and define $T_i = \max(t_{1,i}, \dots, t_{e,i})$, and we then refine equation (4) as follows:

$$\phi_{i,e} \sim \mathbf{N}\left(b_{0,i} + \sum_{k=1}^K b_{k,i} S_k(T_i, t_{e,i}, P_{0i})\right). \quad (20)$$

Alternatively, one might assume that the party with the highest seat shares does not change its position (“winner” hypothesis). In fact, large parties tend to involve more diverse factions, which might hinder policy change (Tsebelis 2002). Positional changes of political parties might also lead to losses of credibility and votes. In the robustness section, we will assess the implications of using the “winner” hypothesis instead of the “incentive” hypothesis.

5 Results

We implement the model in JAGS and R (Plummer 2003; R-Core Team 2012), discard the first ten thousand iterations as burn in, and draw another twenty thousand values.¹² For data storage reasons we only save every tenth draw, yielding a posterior sample of two thousand draws. We run two chains in parallel, generating multiple streams of pseudo-random numbers from the JAGS L’Ecuyer RNG. The Gelman and Rubin (1992) convergence diagnostic supports our choice of run length. In the following, we discuss the posterior distributions of the latent party positions, the factor loadings as well as the bias parameters.

5.1 Party Positions and Factor Loadings

One way to inspect our model is to compare the prior and posterior distributions for the latent party positions, ϕ . For this purpose, we calculate all mean party positions from the posterior distribution, which we call the MCSS. The upper panel of Fig. 1 plots the distributions of the MCSS for each party family, and the lower panel shows the prior densities. Political parties belonging to the same family are located in the corresponding “ideologically delimited space” (Budge 1994, 446). Table 4 lists the prior and posterior means as well as standard deviations for the different party families. Overall, the posterior means are slightly more extreme than the prior means, which indicates that the party family positions are more spread out than the prior information suggests.

To quantify the fit between prior densities and posterior distributions, we take a closer look at the difference between prior densities and posterior distributions of party positions. We quantify the fit by calculating the share of parties in each party family that are estimated to lie in the region between the 2.5% and 97.5% quantile of the prior distribution. The result is listed in column 4, Table 4. Overall, the fit is very high, in particular for the Christian Democrats, the Greens, and the Nationalists, followed by the Liberals, the Left, and the Social Democrats. On average, 93% of the party positions are located in the subspace where the priors expect them to be.

¹² The replication materials are available at <http://hdl.handle.net/1902.1/29312> (König, Marbach, and Osnabrügge 2013).

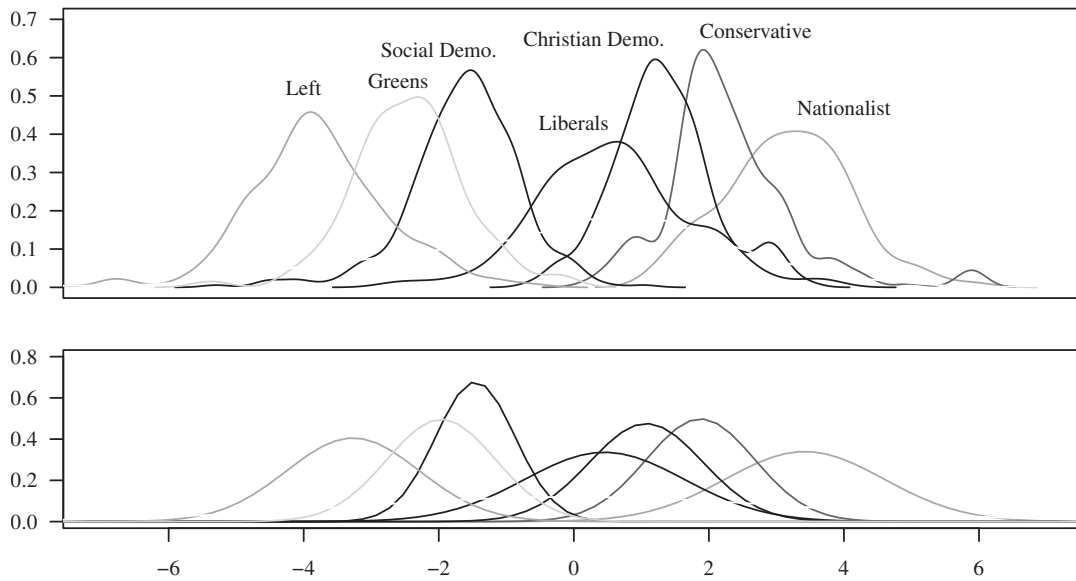


Fig. 1 Party positions: estimated posterior densities of the party position means by party family (upper panel) and the party family prior densities (lower panel).

Table 4 Prior information: means and standard deviations of the prior density for each party family, j , and the means and standard deviations of the posterior party position means for each party family. The last column shows the overall fit, that is, the share of estimated party positions that is in between the 2.5% and 97.5% quantile of the prior density

Party family	Prior		Posterior		Fit rate
	$b_{0,j}$	P_{0j}	Mean (ϕ_j)	sd (ϕ_j)	
Nationalist	3.42	1.17	3.16	0.92	1.00
Conservative	1.88	0.80	2.29	0.96	0.92
Liberals	0.46	1.19	0.58	1.12	0.97
Christian-Demo.	1.07	0.84	1.33	0.75	1.00
Social Demo.	-1.46	0.59	-1.62	0.83	0.97
Left	-3.27	0.98	-3.82	1.10	0.95
Greens	-1.96	0.80	-2.45	0.83	0.99
None	0.00	3.33	-0.18	2.43	0.91

Another method of inspection is to take a closer look at the loadings of each issue on the common left-right dimension. Note that deterministic scaling techniques assume that all issues equally influence the position of a party on the left-right dimension. Table 5 summarizes the posterior densities of the factor loadings indicating how much the latent left-right position loads on one of the sixteen issue indicators. The higher the estimated factor loadings (in absolute terms), the more a specific issue taps the latent left-right dimension.

The highest factor loadings are military, followed by enterprise, freedom, and traditional morality. Overall, the issues that load most are the ones that are conventionally associated with the left-right concept. Indicators such as internationalism, European integration, and multiculturalism are only correlated negligibly with the left-right position. The 95% Bayesian credible intervals (BCIs) only overlap zero for the issue protectionism. From fifteen out of sixteen indicators, the BCI is not overlapping zero, meaning that fifteen out of sixteen indicators load on the latent factor left-right.

Table 5 Factor loadings: posterior density summary of factor loadings, λ , with posterior mean and 95% BCI

<i>Indicator</i>	<i>Post. mean λ [BCI]</i>
Military	0.40 [0.38, 0.42]
Enterprise	0.35 [0.32, 0.38]
Freedom	0.26 [0.24, 0.29]
Traditional morality	0.26 [0.24, 0.29]
Market	0.23 [0.21, 0.25]
Administration	0.21 [0.19, 0.24]
Macroeconomics	0.20 [0.17, 0.22]
National way of life	-0.20 [-0.22, -0.17]
Quality of life	0.16 [0.13, 0.19]
Labor groups	0.16 [0.13, 0.19]
Welfare state	0.14 [0.11, 0.16]
European integration	0.11 [0.08, 0.14]
Multiculturalism	0.10 [0.08, 0.12]
Internationalism	-0.09 [-0.12, -0.06]
Target groups	0.06 [0.03, 0.08]
Protectionism	0.00 [-0.01, 0.02]

5.2 Bias Parameters

Table 6 summarizes the estimates for the country bias parameters, and Fig. 2 visualizes the estimated time bias across twenty-three countries. We exclude Malta and Cyprus from the figure because we can only estimate a single parameter for them. All bias parameters are estimated on the same scale as the party positions, which facilitates the interpretation of their size and direction: A positive (negative) parameter implies that in the absence of the bias correction, parties from this country would appear to be more right (left) than they actually are.

To provide a substantive interpretation of these biases, we propose to compare the bias parameters with half the distance between the modes of the two largest and most important European party families, the Christian Democrats and the Social Democrats. In most countries, the Christian Democrats and the Social Democrats are the dominant parties of each “camp” on the left-right dimension. Hence, we believe that any bias exceeding this distance, which is equal to 1.5, may discredit estimates of party positions because it makes parties appear to be Social Democrats although they are Christian Democrats, and vice versa.

Regarding the country bias, we find a positive bias parameter for thirteen out of twenty-five countries. Accordingly, in these countries parties appear more rightist than they actually are. In most cases, the BCI overlaps zero and is very large, which implies that we would often need more information to pinpoint the direction of the bias. Column four lists the probability that the bias exceeds 1.5, which is half the distance between Christian Democrats and Social Democrats. For five countries, this probability is well above 50%, and for most other countries it is close to 50%. The risk is smallest in Spain (15%) and largest in Estonia and Latvia (98% and 95%). In other words, it is almost certain that Estonian and Latvian parties appear as Christian Democrats although they are actually Social Democrats. In Spain instead, most parties are placed in the appropriate subspace of the space.

Figure 2 illustrates that the bias varies considerably over time and across countries. In some countries, such as Denmark, the time bias scatters around zero, whereas it exhibits a persistent trend toward one extreme in others, such as in the United Kingdom. This means that the British left-right scale has tremendously changed over time. In longitudinal studies of party systems, the existence of such a persistent trend is especially worrisome since the error is systematic and not random. The average probability for any time bias to exceed 1.5 across all countries and elections is 53%. In the United Kingdom, this probability is rising over the course of the last elections,

Table 6 Country bias parameter: posterior density summary of the country bias parameter, ρ , with posterior mean and 95% BCI

<i>Indicator</i>	<i>Post. mean ρ [BCI]</i>	<i>$p(\rho \geq 1.5)$</i>
Estonia	4.01 [1.63, 6.46]	0.98
Latvia	3.22 [1.26, 5.33]	0.95
Cyprus	2.51 [-0.28, 5.20]	0.77
Lithuania	1.94 [0.05, 3.82]	0.66
Malta	1.94 [-1.82, 5.69]	0.63
Poland	1.49 [-0.11, 3.07]	0.48
Sweden	1.37 [-0.67, 3.44]	0.45
Netherlands	1.21 [-0.72, 3.18]	0.38
France	1.05 [-0.97, 3.00]	0.34
Germany	0.62 [-1.11, 2.32]	0.19
Finland	0.50 [-1.67, 2.69]	0.23
Czech Republic	0.10 [-3.09, 3.13]	0.35
Ireland	0.09 [-2.17, 2.31]	0.20
United Kingdom	-0.07 [-2.47, 2.35]	0.24
Spain	-0.12 [-2.05, 1.82]	0.15
Hungary	-0.45 [-2.36, 1.48]	0.17
Austria	-0.54 [-2.92, 1.93]	0.26
Slovenia	-0.57 [-2.65, 1.51]	0.23
Belgium	-0.64 [-3.97, 2.41]	0.36
Greece	-0.69 [-3.16, 1.81]	0.30
Denmark	-0.84 [-2.85, 1.23]	0.29
Italy	-0.93 [-3.11, 1.14]	0.31
Luxembourg	-1.13 [-4.18, 1.89]	0.47
Portugal	-1.21 [-3.57, 1.23]	0.42
Slovakia	-1.48 [-4.22, 1.06]	0.48

approaching 92% in the 2010 election. In Denmark instead, the probability is not larger than 65% (1977) and has fallen in the last two elections to 36% (2007).

These results suggest that country and time effects might seriously bias estimates of party positions on the left-right scale from coded manifesto data. Our findings on country effects corroborate research on party systems in transformation countries, where it remains difficult to distinguish between Christian and Social Democrats without correction. Similarly, the recent development in the United Kingdom suggests that the current location of the political parties is difficult to compare with the position of the same parties some years ago. Both findings may have important ramifications for our understanding of the interaction between parties and institutions in comparative politics. Whether and to what extent they may change our perspective will be shown by comparing the MCSS with existing estimates.

6 Validity and Robustness

Similar to Ho and Quinn (2008), we evaluate our estimated party positions using several robustness criteria. First, we show how far the MCSS are correlated with existing measures of party positions (convergent validity or cross-validity). We find that the MCSS diverge more from previous scores based on manifesto data than from expert survey data. Second, we check their robustness by changing some of our key assumptions, vary the sample composition, and the source of the prior information. The results indicate high robustness of the MCSS against any changes of the key assumptions, sample composition, and source of prior information.

In addition to these overall robustness checks, we also discuss the construct validity of our estimated party positions at the country level. For this purpose, we evaluate to what extent our MCSS are consistent with existing qualitative knowledge of the left-right positions in two

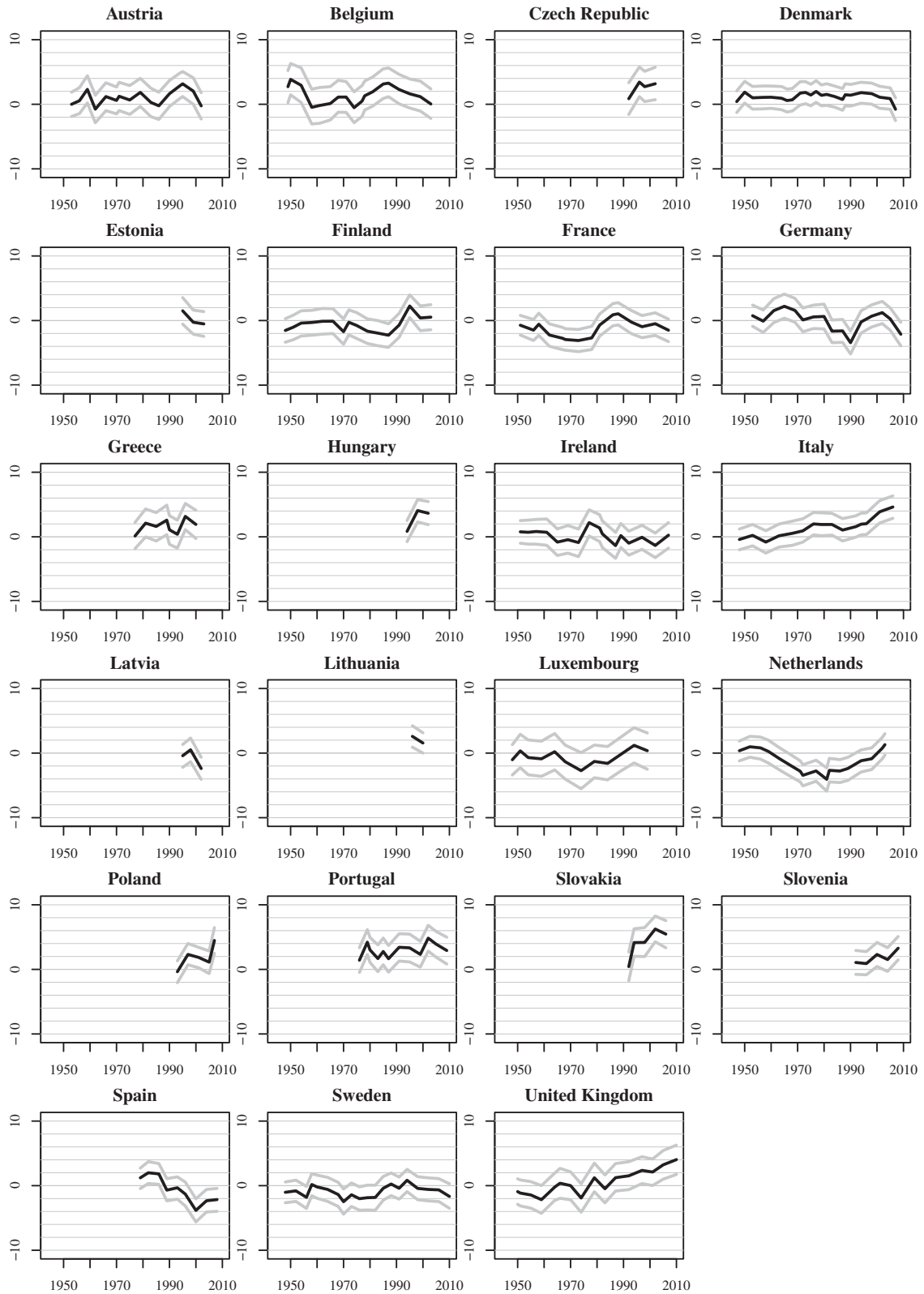


Fig. 2 Time bias parameter: posterior density summary of the time bias parameter, θ , and 95% BCI. For this figure we only included countries where we observed at least three national elections.

countries, France and Sweden.¹³ As stated by Dinas and Gemenis (2010, 428), existing studies often disregard this step and do not check how far the estimated party positions are plausible at this level of analysis.

6.1 *Convergent Validity*

In order to assess the convergent validity, we compare our MCSS with three prominent left-right scales: the vanilla scores, Rile, and expert data. The vanilla scores are estimated by running a standard factor analysis on the 56 CMP categories (Gabel and Huber 2000). Rile is calculated by subtracting the sum of several categories associated with left political ideas from the sum of several categories referring to ideas from the political right (Laver and Budge 1992). The expert data stem from Hooghe et al. (2010) and Steenbergen and Marks (2007).¹⁴

Figure 3 shows the results. The first two panels plot the MCSS against the vanilla and Rile scores. Both correlation coefficients are approximately 0.6. Vanilla and Rile scores tend to place some parties much more to the extreme than the MCSS. The third panel compares the MCSS for parties competing in the elections of 1999, 2002, and 2006 with expert survey data from the same years (Steenbergen and Marks 2007; Hooghe et al. 2010). The correlation here is quite high, $r = 0.86$. In fact, it is the highest among all four score sets in a cross-correlation matrix.¹⁵ This suggests that our scores differ substantially from alternative estimates derived from the manifesto data but are more in line with expert survey data. This difference is even more pronounced when we apply the different estimates to important theoretical concepts. We calculated the median party position and the polarization coefficient on the basis of our model and the vanilla method. In a linear regression of the median party positions based on our scores on the median party positions derived from the vanilla scores, we find little correspondence. The R^2 is only 0.01.¹⁶ Similarly, a linear regression with the polarization coefficients only yields a modestly larger correspondence $R^2 = 0.18$.¹⁷

6.2 *Robustness*

To analyze how far the MCSS depend on our assumptions, we estimate our latent variable model under the same conditions but changing some of the key assumptions. Specifically, we run the model without Legendre polynomials, bias parameters, and informed priors. We also test to what extent the source of the prior information matters and whether the composition of the sample impacts our results. The Gelman and Rubin (1992) convergence diagnostic again supports our choice of run length for each estimated model.

To replace our priors from expert surveys, we start with running a model where we used Eurobarometer surveys spanning the period 1970–2002 to calculate the informed prior distribution on position parameter (Schmitt et al. 2008). The Eurobarometer is a public opinion survey across the EU member states. We used the questions on the respondents' left-right self-placement and their vote intention to infer a party family's position. The correlation between the MCSS and the alternative scores from this model is very high ($r = 0.98$).¹⁸

¹³ Although we only discuss France and Sweden here in depth, readers are referred to the supplementary appendix for similar figures of all other 23 countries.

¹⁴ Using the same data source for the cross-validation as for the prior specification is not tautological in a sense that the high correlation follows directly from our usage of priors derived from the same expert survey. The priors only reflect information about the position of the party family in the left-right space, not about the parties' positions themselves.

¹⁵ The vanilla and Rile scores are correlated by $r = 0.47$. The correlation coefficient between Rile and expert survey scores is $r = 0.71$; vanilla and expert survey scores are correlated by $r = 0.57$.

¹⁶ Intercept: 7.04(0.05), Coefficient: 0.09(0.05).

¹⁷ Intercept: 1.49(0.20), Coefficient: 0.23(0.03).

¹⁸ The large amount of survey responses in the Eurobarometer also allows us to check whether the party families' means and variance change over the years, which could have an impact on our posterior estimates. We find that the annual differences are not significant. In other words, the party families' means and variances in 1983 are indistinguishable to the data in 2002.

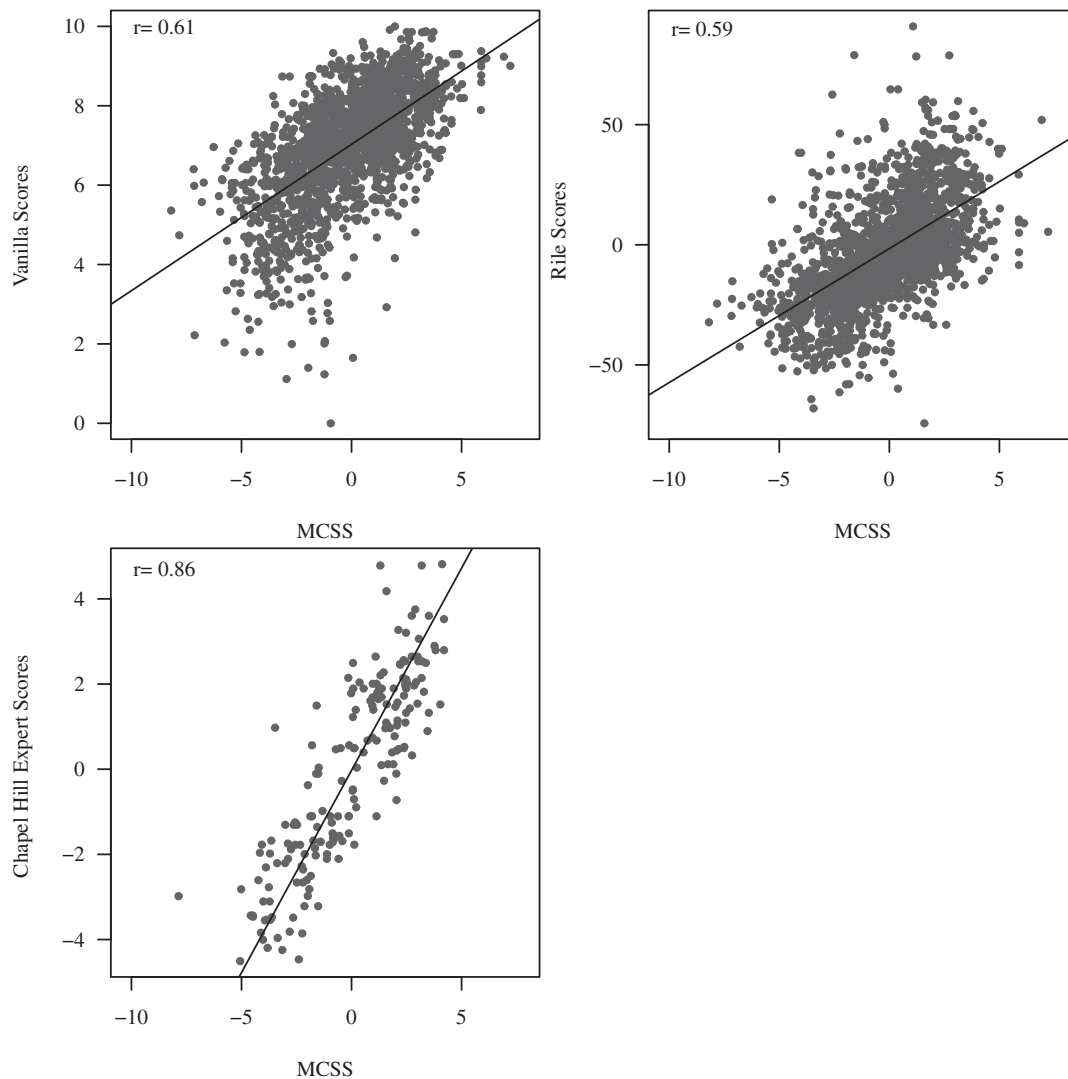


Fig. 3 Convergent validity: correlation between different left-right estimates derived from manifesto data (vanilla, Rile, and Chapel Hill expert scores) and the MCSS.

Next, we drop the Legendre polynomials from our model assuming independence between positions of the same party over elections. We find that the estimated party positions from this model are correlated with the MCSS by $r = 0.81$. This means that the Legendre polynomials have some influence on our estimates. Next, we additionally drop all bias parameters. The resulting model is equivalent to a simple Bayesian factor analytical model with informed prior distributions on the latent factors. The correlation between the alternative scores from this model and our MCSS is relatively high ($r = 0.92$). Finally, we also drop the informed prior distributions on the positional parameters. Technically, we identify the parameters of this model by fixing the position of the German Christian Democrats (CDU) at 2 and the German Social Democrats (SPD) at -2 for the 2005 election. Note that this identification constraint does not influence our results. The correlation coefficient between the MCSS and the scores from this alternative model is $r = 0.84$. These results are very encouraging since they demonstrate that changing most of our key assumptions hardly impacts our MCSS scores. What remains is the question on the influence of our bridges on the estimates.

Regarding our bridges, we examine to what extent our MCSS are conditional on the “incentive” and “zero hour” hypotheses. First, we run an alternative model which keeps constant the position

of the party winning most seats in absolute terms (“winner” hypothesis). The correlation between these estimates and our original ones is $r = 0.77$. Second, we use alternative country bridges by assuming that the position of the national party is equal to the one of the national delegation to the EP when the national and European elections are closest to each other (“closeness” hypothesis). The correlation coefficient is $r = 0.78$. This again suggests relatively high robustness of our original estimates when using alternative bridge observations.

Finally, to provide some evidence that the choice of the sample composition does not determine the results, we estimate a model where we only include data from 1975 onward. This cutoff is convenient since it allows us to retain all cross-country bridges and, hence, attribute any changes of our estimates to the sample composition. We find that the results are almost identical ($r = 0.99$). In addition, the posterior densities of the factor loadings also do not change much. One interpretation of this finding is that the meaning of the latent left-right dimension is relatively constant over time. If all data points before 1975 are dropped, the difference distribution of the posterior loading distributions is not overlapping zero in only three cases: quality of life, multiculturalism, and target groups. And even in these three cases, the difference is marginal (the posterior mean difference average is 0.06, 0.06, and 0.04). This empirical result provides some evidence that our assumption about fixed factor loadings is reasonable.

In sum, our scores are robust to changes in key assumptions, the sources of prior information, and the composition of the sample. Compared to the polynomials and the priors, the bias parameters have a smaller impact on the estimates if we drop them from the model. However, when using alternative bridges, their impact becomes larger. Nevertheless, as our previous discussion has shown, the estimated bias parameters are substantial and systematic, which is why scholars should correct for those biases in their analyses.

6.3 Construct Validity

Next, we analyze the construct validity of our estimates in two countries: France and Sweden. We selected France and Sweden for two reasons. In both cases, our estimates differ considerably from existing measures, which allows us to illustrate in more detail in how far this divergence is plausible. Furthermore, the party systems of Sweden and France are very different (Sartori 1976). While the Swedish party system is often classified as a consensus type, the French party system is considered polarized. We briefly discuss the main characteristic of the French and Swedish party systems before we compare our estimates with the vanilla scores.¹⁹

The main characteristic of the French party system is the bipolarity between left- and right-wing party camps (Lowe 1989; Borella 1990), which is caused by the French electoral system. A typical feature of this polarized camp pattern is that there is no center party that might be in a position to change majorities by switching its support from a left- to right-wing party camp. French bipolarization was especially strong in the Fourth Republic (1946–58) due to antisystem parties. However, Knapp and Wright (2006, 256–59) suggest that this polarization is decreasing since the elections in 1981. This is especially true for the period after 1997 when the Communists and Greens became junior parties of government.

Sweden is considered as a typical example for a frozen party system due to its stability and continuity (Vedung 1988; Foldal 1989). This is anchored in the relatively constant cleavage structure and the consensual nature of the political system (Lipset and Rokkan 1967). However, the five-party system changed in the 1988 election when the Greens entered parliament and in 1991 when the Christian Democrats also entered parliament. Another characteristic of the Swedish party system is the dominance of the Social Democratic Party, which was in government continuously from 1932 to 1976 and sometimes supported from right-wing parties due to their own disunity (Jahn 2009, 125).

¹⁹ The supplementary appendix includes a graph on the Rile scores. We do not discuss it separately here since it is affected by the same problems as the vanilla scores.

The upper panels of Figs. 4 and 5 plot the MCSS with a 95% BCI for selected parties, and the lower panels show the vanilla scores for the two countries. The BCI resembles a frequentist confidence interval, but has a different interpretation. A BCI refers to the interval where the parameter lies with a probability of 95%. Instead, the 95% confidence interval indicates where a parameter lies in 95% of the hypothetical replications of the analysis. Note that whereas overlapping confidence intervals imply that the two estimates are statistically indistinguishable, overlapping BICs merely mean that there is some (computable) probability that the parties favor the same position.

The MCSS suggest that parties change their left-right positions over time but their overall ordering on the left-right dimension remains stable. This general pattern is in line with the literature arguing that party systems are relatively stable and change only gradually over time (Lipset and Rokkan 1967; Bartolini and Mair 1990; Hanley 1999; Mair 1999; Mair 2001). The vanilla scores in turn are subject to a puzzling extreme volatility with large zigzag movements of parties. In some instance, Swedish parties jump through a quarter of the policy space from one election to the next (e.g., the Communists), leapfrog (Social Democrats with the People's Party), and shift in parallel from over elections (e.g., 1990). Furthermore, the vanilla scores indicate that in the 2007 Swedish parliamentary elections parties suddenly converged almost entirely to a single point, which is certainly not in line with the political reality.

For France, most surprising are the scores for the right-wing Front National, which vanilla locates in the 1990s on the same position as UDF and UMP. In 2007, the party is even put in the middle of the political spectrum, close to the Greens and Socialists. In this election, Front National advocated the right-wing policy “Les Français d’abord” (French First), which includes the expulsion of immigrants and more recently the prohibition of new mosques. The MCSS in turn

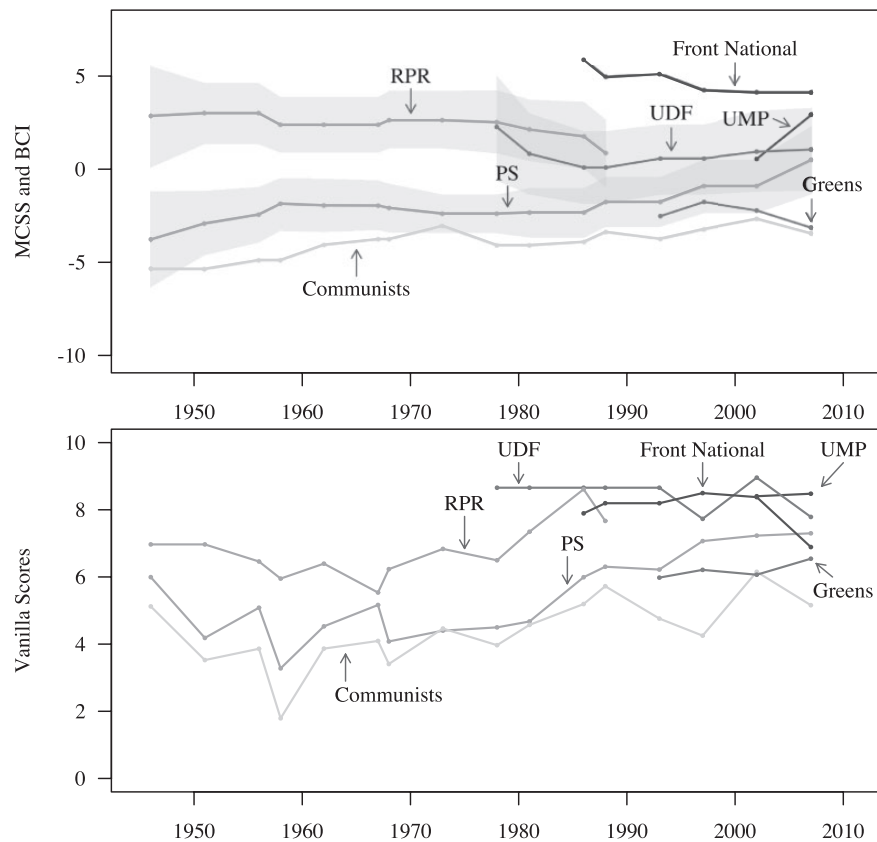


Fig. 4 French party system: MCSS and 95% BCI for selected parties, and vanilla scores for seven French parties.

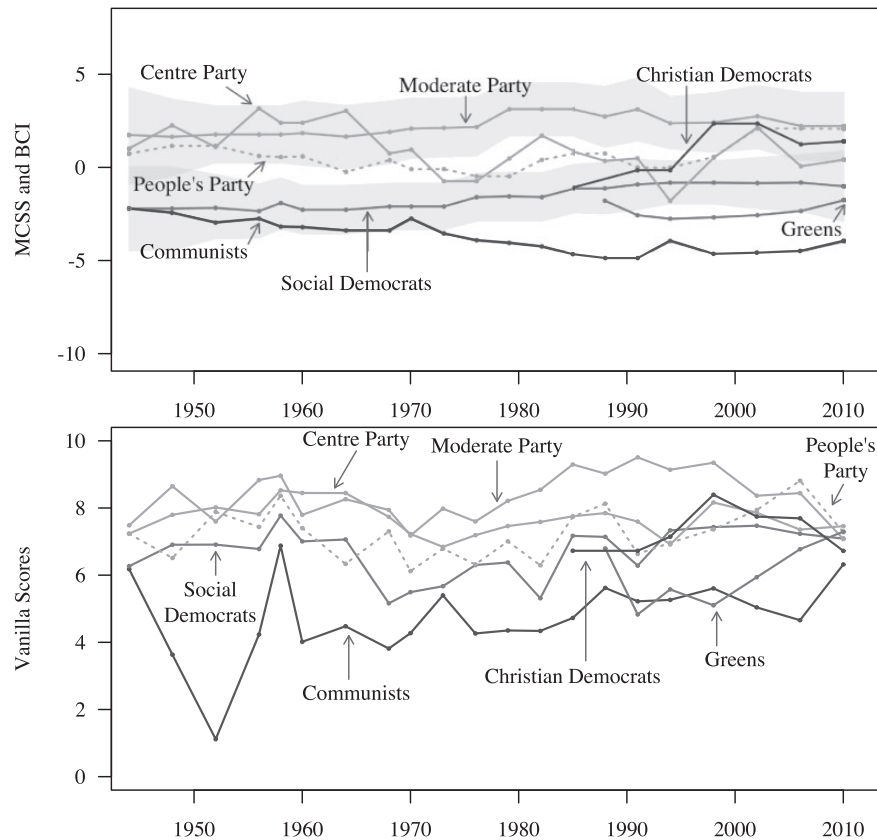


Fig. 5 Swedish party system: MCSS and 95% BCI for selected parties, and vanilla scores for seven Swedish parties.

place the Front National at the far right of the political spectrum, being very different from the UDF. In fact, the average posterior distance between these two parties is 3.2 left-right points in 2002 and 2007. The posterior probability that this distance is smaller than zero, that is, that the UDF is actually to the right of Front National, is 1.65% (for the same two elections). Note that this pattern remains when we exchange our key assumptions.

In our view, the MCSS appear to be more consistent with existing knowledge about the French and Swedish party systems. The inspection reveals that existing estimates overestimate the positional changes of political parties and misplace some parties. This country-specific evidence applies to many further countries and the development of their party systems, which we cannot present in more detail here. Our application indicates that previous estimation methods have difficulties in locating political parties from different party systems into a common space.

7 Discussion

This article introduces a new method for estimating party positions of political parties across countries and time by introducing a latent variable model for manifesto data. Our estimation strategy builds upon the logic of recent methodological advances on placing actors in a common space (Bailey 2007; Shor and McCarty 2011; Gschwend, Lo, and Proksch 2012). A general contribution lies in developing a statistical model that addresses country- and time-specific effects. Existing estimates of party positions do not take this potential bias into account (Gabel and Huber 2000; Laver, Benoit, and Garry 2003; Slapin and Proksch 2008; Lowe et al. 2011). We apply the model to estimate party positions from coded manifesto data in Europe and identify bridge observations to simultaneously model country- and time-specific bias parameters for a large set of political parties over a long period of time.

Our statistical model rests on three major assumptions: (1) the linear functional form that relates the latent party positions in an orthogonal space to the observed issue positions which are distributed normally; (2) the dependency between party positions across elections is represented with a set of Legendre polynomials; (3) the time and countries bias is captured by linear, additive bias parameters. When we apply the model, we make three additional assumptions: (4) to identify the country-specific bias parameters, we argue that political parties participating the first time in a EP election take the same position as in the previous national election (“zero hour” hypothesis); (5) for the time-specific bias parameter we assume that the political party with the largest relative gains in seat shares has no incentive to change its position vis-à-vis other parties in the next election (“incentive” hypothesis); (6) we assume that parties belonging to the same party family take similar positions in specific segments of the left-right dimension, and we locate these segments using expert survey data.²⁰

We discussed alternative assumptions and checked the robustness of our findings on the aggregated and country level. Overall, our robustness checks indicate similar findings when we drop or change the above-mentioned key assumptions. We showed that our method generates explicitly comparable estimates of positions of a large number of political parties. These scores offer three particular advantages over previous estimates: (1) they are comparable across countries and time periods; (2) our scores are robust to changes in the empirical estimation strategy, which means that changes in the assumptions only modestly affect our findings; (3) our scores exhibit a high convergence validity with expert survey data and a high construct validity. We demonstrated the application of our method by estimating the positions of 388 parties competing in 238 elections across twenty-five countries and over sixty years. Our results indicate that existing measures are likely to be affected by country- and time-specific biases. This finding is not of surprise when we acknowledge that political parties compete in elections within specific countries at a particular point in time. The bias parameters suggest that not controlling for country- and time-specific effects produces misleading estimates of party positions.

We believe that our findings may have several implications for studies in comparative politics: First, the size of the estimated bias parameters can be used to guide researchers’ case selection in studies using manifesto data. When researchers select their cases accordingly, they do not necessarily have to use a statistical model to correct for the time and country bias in the manifesto data. For example, our findings suggest that comparing party positions from the United Kingdom and Ireland would not involve a serious country bias since the corresponding parameters are almost zero. Similar, the estimated time bias parameters indicate that comparisons of Danish parties over time are relatively unproblematic (the time bias is on average zero). Second, when scholars use our scores to test theories in comparative politics, they can be more confident about their results. Since our MCSS are comparable across countries and time periods, their statistical test will be less biased by systematic measurement error arising from country- or time-specific effects. Furthermore, scholars can also propagate the uncertainty about a party’s position into their analysis by leveraging the full posterior distribution.

Our method can serve as a basis for future methodological work. Our findings underline the importance of assessing the comparability of scales in cross-country and longitudinal studies. One way to advance our method would be to extend our statistical model by representing the full data-generating process while not compromising the tractability of the estimation. Another extension would be to adapt existing computer-based models based on the distribution of words to take into account country- and time-specific effects. Advances along these lines will provide an opportunity to make progress in the study of parties, institutions, and voters in comparative politics.

²⁰ Some readers might object that our assumptions to estimate the model are too strong and that it is not worth it to trade off the empirical uncertainty about the comparability of the data against the theoretical uncertainty about the validity of our model assumptions. Although we agree with these potential critiques about the existence of this trade-off, we stress that we find it more useful to have clearly specified uncertainty about some underlying modeling assumptions than uncertainty buried in the data.

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