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Do Individual Heterogeneity and Spatial Correlation Matter?

An Innovative Approach to the Characterisation of the European Political Space.*

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In this paper we refine the interpretation of the European two-dimensional political space and the investigation of its determinants compared to the approach commonly adopted in the spatial voting literature. Specifically, we take into account heterogeneity and cross-correlation among legislators by explicitly including into the model a spatial effect which, in turn, relies on new sets of linguistic, geographical, institutional and cultural metrics. We confirm that the first dimension of the European political space is mainly explained by the Members of European Parliament's ideological position on a left-right scale. We also find that correlation across legislators plays a significant role in explaining the first dimension when their pairwise distance is defined according to an individualism index, which turns out to be closely related to left-right ideology positioning. Even more interestingly, we show that "space" intended in a broad economic sense plays an important role in interpreting the second dimension of the political spectrum. The most relevant metric that induces spatial effects along the second dimension is based on an institutional index. Moreover, we also find that the second dimension is influenced by the gender composition of the political parties.

Keywords: European political space, spatial autoregressions, NOMINATE, proximity matrices, economic distances.

JEL codes: D72, C21.

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1. Introduction

The European Parliament is an institution of particular interest from the point of view of economists and political scientists. It is a relatively young institution, which has increasingly gained more power in several aspects, such as the set of issues that it is called to decide upon, the number of voters represented, and the number of votes it casts. A peculiar characteristic of the European Parliament is the strong heterogeneity of its components. Members of the European Parliament (MEPs) are elected in districts that do not cross national borders, from lists chosen by national parties, and with electoral rules that, albeit proportional, are country-specific. Therefore, they represent their countries and their national parties, as well as the European Political Group they belong to. Moreover, they are only accountable to their national electorate and with rules that differ from country to country. As a consequence, politics in the European Parliament are likely to be subject to more influences and effects than politics in national parliament, as Hix, Noury and Roland (2007) discuss extensively.

The understanding of what drives legislators' behaviour in the European Parliament is of particular relevance to evaluate for example the effect of policy changes, such as changes in the electoral rules of the European Parliament, or institutional changes, such as changes in the composition of the Parliament due to an enlargement of the EU, or to phenomena such as Brexit. Cross-influences among legislators may change the consequences of policies or institutions enhancing or weakening their effects, thus favouring or opposing the policy makers in reaching their objectives.

The sources of heterogeneity among MEPs mentioned above are also sources of potential spatial correlation across legislators, where space is intended in a broad way that includes economics/cultural characteristics. The aim of this paper is to investigate whether these correlations can influence the positioning of legislators in a policy space which is intrinsically multidimensional. Like previous literature, we focus on the determinants of EU legislators' behaviour by the use of the NOMINATE scaling method applied to roll call votes. The NOMINATE method delivers a characterisation of the dimensions of the policy space, that is, the number of evaluative characteristics that influence the policy decisions, and ideal point estimates on each policy dimension for each MEP. It has been shown in the literature (Hix, Noury and Roland, 2006) that the policy space is described carefully by to two policy dimensions. The substantive meaning of the dimensions of the policy space needs then to be interpreted with the use of additional exogenous information. The novelty of our approach compared to that currently adopted in the literature (e.g. Hix, Noury and Roland, 2006, and references therein) is that the interpretation of this substantive meaning is performed introducing a spatial component in the regressions, by means of the spatial autoregression models (SARs, henceforth). SAR models have been applied to political science and political economy in the investigation, for instance, of international relations, trade and the effect of public policy (e.g. Gleditsh, 2002; Cho, 2003; Lacombe, 2004; Beardsley et al., 2006). However, even though the analysis of voting patterns seems a natural field of application of spatial econometrics models, to the best of our knowledge such techniques have never been applied in this context.

SARs are flexible models that take into account potential spatial links across observations. The notion of spatial correlation is embedded into a set of *weights* which define the shape of the pairwise interactions between observations. In turn, such weights are exogenously constructed from suitable economic distances that include the standard geographical proximity as a special case. In this paper, we first introduce a set of metrics, which capture geographical, linguistic, institutional and cultural distances. We then analyse whether the correlations described by those metrics are possible channels through which legislators affect each other. We find that each single legislator's position in the policy space does influence the position of other legislators. Specifically, the evidence suggests that spatial correlations matter in both policy dimensions. The first dimension is mostly explained by the positioning on the left-right ideological scale of each national party, like in the literature. We do however observe a significant effect of the spatial correlations based on the Individualism index by Hofstede, Hofstede and Minkov (2010). We notice that Kallio and Niemelä (2014) show that the ideological left-right orientation is correlated with attitudes towards individualism. Therefore, we highlight how the effect of the ideological left-right positioning on the first dimension is both direct, like in the literature, and indirect, through this spatial channel. The second dimension has instead an ambiguous characterisation in the literature. We confirm the relevance of the standard exogenous explanatory variables. Additionally, we observe two significant spatial effects. The first one is the effect of the institutional proximity based on the Parliamentary Power Index by Fish and Kroenig (2009), which is strong and robust. The second is the effect of the cultural proximity based on the Masculinity index by Hofstede, Hofstede and Minkov (2010), which is weaker and less robust. As the Masculinity index is correlated with the presence of women in politics, we also include the gender composition of each national party among the covariates. Our gender variable is strongly significant in explaining the second dimension, but the spatial effect of the Masculinity index remains.

Over the past few decades, a growing literature in political economy has focused on the determinants of legislators' behaviour in Congress by the use of roll call methods (Poole and Rosenthal, 1997, 2011; Rosenthal and Voeten, 2004; Carey, 2003). The method of roll-call voting data scaling (Poole, 2005; Poole and Rosenthal, 1985, and references therein) is based on data sets containing every individual vote on every roll call. Recently, these methods were interestingly used to analyse supranational settings such as the European Parliament (Hix, Noury and Roland, 2006; Hooghe, Marks and Wilson, 2002) investigating the dimensionality of the European Parliament policy space.

Hix, Noury and Roland (2006) analyse this issue by aggregating observations at the national party level, and show how the first dimension can be viewed as related to the national party ideological positioning on a left-right scale, while the second one is related to the position about European integration. Unlike Hix, Noury and Roland (2006) we explicitly take into account possible spatial correlation across legislators' positions by means of SARs models. We confirm their interpretation of the first dimension as driven by the ideological positioning on the left-right scale. We provide further insights on the second dimension, highlighting the presence of spatial effects, and the effect of gender composition of the national parties. Theory and applications of spatial econometrics, on the other hand, have increased significantly over the last few decades, as theorists and practitioners have become increasingly more aware that lack of cross-sectional correlation is not a realistic assumption in most empirical settings. For an exhaustive review of spatial models, their taxonomy and their peculiarities we refer to Anselin (1988), while an up-to-date survey of the advances in inference techniques for spatial data is given in Elhorst (2014).

Section 2 describes the methodology, Section 3 introduces the proximity matrices, Section 4 presents the results, Section 5 discusses the robustness checks and Section 6 concludes.

2. Methodology

In order to describe and analyse the European policy space, we rely on the methodology known as NOMINATE (Poole and Rosenthal, 1997, and references therein), which in a nutshell is a scaling method for roll call votes data. These are approximately one third of all votes that take place in the European Parliament. Roll call votes are obligatory for the final vote on legislation, while for other votes the roll call procedure may be requested by an European Party Group or by a group of at least 40 MEPs. The NOMINATE technique has been designed to estimate the position of each legislator's bliss points in the policy space, starting from a random utility function which is known only up to a finite set of parameters. The first step of NOMINATE is thus a multi-dimensional scaling procedure, followed by the estimation of unknown parameters of the utility function by maximum likelihood. The steps are then iterated until convergence. It is worth pointing out that the multi-dimensional scaling (and thus the whole estimation procedure) relies on a criterion of similarity across legislators. The pairwise index of similarity between legislators is based on the number of times they vote in the same way in roll calls. Hence, the outcome of NOMINATE are estimates of relative positions across legislators rather than of their actual ideal points (which are only known up to a scale factor) along a number of orthogonal dimensions of the political space that has to be postulated *ex ante* by the practitioner. Hix, Noury and Roland (2006) show that estimates of legislators' bliss points along a two-dimensional spectrum correctly predict about 90% of roll call votes, so that the European policy space can be considered two-dimensional. Thus, following Hix, Noury and Roland (2006), in this setting the NOMINATE procedure returns estimates of legislators' ideal points along two dimensions. The NOMINATE procedure has been extended to allow for elliptic, rather than spherical, indifference curves associated to the utility function of legislators (the so-called weighted NOMINATE) and for a dynamic setting, so that several consecutive legislatures can be analysed at the same time (e.g., Poole and Rosenthal, 2001, and references therein).[1]

As NOMINATE and its variants deliver estimates solely based on a similarity index between each pair of legislators, its main shortcoming is the lack of information provided about the economic and political meaning of the dimensions. The substantive meaning of each dimension can be investigated using regression analysis (e.g. in Hix, Noury and Roland, 2006).

The novelty of our approach compared to that currently adopted in the literature (e.g. Hix, Noury and Roland, 2006, and references therein) is the introduction of a spatial component in the regressions, by means of the well known spatial autoregression models (SARs, henceforth).[2] In many empirical problems SARs offer a useful framework for describing data which are generally irregularly spaced, without a natural ordering and/or a geographical interpretation, such as legislators' coordinates. In SAR models the notion of possible irregular spacing, based on general economic distances, is embodied in an $n \times n$ weight matrix (n being sample size), denoted W, which needs to be chosen by the practitioner. In general, the economic distance between legislators i and j is defined as the distance between u_i and u_j, where u_i and u_j are vectors of characteristics pertaining to legislators i and j, respectively. The distance between u_i and u_j might be defined in an Euclidean sense. A vast choice of relevant economic distances among legislators is discussed in Section 3. Let w_{ij} be the $(i, j)-$th element of W. Conventionally, $w_{ii} = 0$ for $i = 1, ...n$, i.e. the spatial interaction of each legislator with itself is set to zero. Often, but not exclusively, w_{ij} is defined in terms of the inverse of an economic distance between units i and j. In other cases, as legislators belong to different regions or countries, W can be chosen according to a contiguity criterion, i.e. $w_{ij} = 1$ if their regions or countries share a border and $w_{ij} = 0$ otherwise. In our empirical analysis we also normalise W so that the entries in each row sum to one.

Let y and X be standard sets of observable variables, indicating respectively dependent and independent variables, while ϵ indicates a vector of independent and identically distributed (iid) normal random variables, with mean zero and unknown variance σ^2.

[1] Our results are derived by applying the weighted dynamic version of NOMINATE, DW-NOMINATE, to take full advantage of a larger dataset. As we only deal with five legislatures, the standard static model would probably deliver similar results.

[2] For exhaustive surveys of spatial models and applications see for instance Anselin (1988) and Arbia (2006).

The standard SAR is defined as

$$y = \lambda W y + X\beta + \epsilon, \tag{1}$$

for some unknown parameters β and a scalar unknown parameter λ. Specifically, the significance and magnitude of the estimate of λ define the spatial effects. According to eq. (1), the dependent variable of each unit is not only explained by its own vector of characteristics, but it is also related to a weighted average of the dependent variables of neighbouring units.

The model of eq. (1) is a very parsimonious method of describing spatial dependence, conveniently based only on economic distances rather than actual locations, which may be unknown or not relevant. Although a major drawback of SAR models is the ex ante specification of W, to which parameter estimates are sensitive, eq. (1) has been widely used in practical applications because of its flexibility. The possibility of considering several specifications of W allows us to investigate the effects of multiple sources of interactions among legislators.

In our work, we also consider the slightly more general version of eq.(1) known as spatial Durbin model, defined as

$$y = \lambda W y + X\beta + WX\gamma + \epsilon, \tag{2}$$

where, in addition to the endogenous spatial component Wy, a direct exogenous interaction effect WX is included.

3. Selection of economic proximities

The main focus of this paper is to understand which correlations among legislators may help explain their positioning on the European policy space. We generate several versions of the proximity matrices so as to assess whether clustering among legislators is influenced by correlations across several national characteristics. We define matrices based on geographical, linguistic and institutional distances, which we call *geopolitical proximities*, and matrices based on the cultural indexes by Hofstede, Hofstede and Minkov (2010) (*cultural proximities*).

Each choice of the proximity matrix W is based on a different distance across countries, and it is built with $w_{ij} = \frac{1}{Distance_{ij}}$. A technical issue to consider is how to set the W entries between legislators with the same nationality. Since the present work aims to shed light on the implications of trans-national correlation across legislators, a sensible modelling choice is to explain each legislator's position in the political spectrum by a set of their own characteristics (possibly country-specific) and by their respective interactions with legislators belonging to neighbouring countries. Thus, we set w_{ij} equal to zero if legislators i and j belong to different countries.

Geographical proximity. We first start with a characterisation of geographical proximity based on the distance in kilometres between capitals of European member states of legislators i and j, measured as the average of the shortest outbound and inbound routes suggested by Google Maps.[3]

Linguistic proximity. The second proximity measure is based on a linguistic metric. We measure the distance between legislators based on their home country languages. For a comprehensive analysis of distances across languages we refer to Ginsburgh and Weber (2011). As they show, linguistic proximity has an effect on economic and political outcomes such as trade, immigration and voting behaviour. We build a linguistic proximity matrix based on the lexicostatistical distance by Dyen, Kruskal and Black (1992).[4] Lexicostatistical distances are based on the vocabulary of a language. They are built from the computation of the percentage of words which share a common origin, defined by linguists as cognate words (such as the English *father* and the German *Vater*), in a set of common "list of meanings".

Institutional proximity. The third choice of proximity measure is related to the MEPs' institutional background. To characterise the institutional environment that is familiar to each MEP we consider the score of his home country in terms of the Parliamentary Power Index by Fish and Kroenig (2009). The authors, in their Legislative Power Survey, identify 32 possible powers that a legislature may have (e.g., power to appoint the prime minister or the chairman of the central bank, power to grant pardons or amnesties, immunity from dissolution in case of dissolution of the government) and compute the Parliamentary Power Index as the fraction of such powers that a legislature has. Using the PPI we create a distance between legislators' home countries, defined as $|PPI_i - PPI_j|$.

Cultural proximities. Finally, we consider a set of proximities based on the six cultural indexes by Hofstede, Hofstede and Minkov (2010), which describe the attitudes of national cultures towards different issues that may influence legislative decision making. The six indexes are:
Power Distance Index. The PDI index measures the extent to which less powerful members of institutions expect and accept unequal distribution of powers. High PDI

[3]For robustness, we also considered a second geographical matrix, where the distance between capitals is measured in terms of flight duration. The two distances are highly correlated and lead to the same considerations. Results are available from the authors under request.

[4]For the construction of the linguistic matrix, French Belgium and Flemish Belgium were considered as separate countries. The DKB lexicostatistical distance is not available for pairs which involve legislators from Finland, as their official language is not Indo-European. We set all these values to 0 (which corresponds to minimal proximity, or maximal distance).

scores are correlated with a political spectrum with a weak center and strong right and left wings, and fewer parties.

Individualism Index. The IDV index classifies societies based on whether they display individualism (preference for a social framework in which individuals take care only of themselves and their close family) or collectivism (preference for a framework in which individuals expect their relatives or members of a particular ingroup to look after them). Therefore high IDV scores are correlated with societies where privacy and individual freedom prevail over collective interests.

Masculinity Index. The MAS index classifies societies based on the distinction (or absence of distinction) of emotional roles by gender. High MAS scores are correlated with preferences for large organizations (vs. small), with the tendency of resolving conflicts by letting the strongest win (vs. negotiation) and with low participation of women in politics and management.

Uncertainty Avoidance Index. The UAI index measures the extent to which members of a culture feel threatened by ambiguous or unknown situations. High UAI scores are correlated with the presence of many and precise laws, with a slow judiciary process and with a low participation in politics.

Long-Term Orientation Index. The LTO index measures the weight that societies give to virtues oriented towards future rewards (such as perseverance) as opposed to virtues related to the past and the present (such as respect for tradition). LTO scores are correlated with investment choices, nationalism and fundamentalism.

Indulgence vs. Restraint Index. The IVR index measures whether a culture as a tendency to allow relatively free gratification as opposed to the conviction that such gratification needs to be regulated by strict social norms. IVR scores are correlated with the importance of freedom of speech, the importance of maintaining order and the number of police officers.

We define the distance between countries based on Index k as $|Index_i^k - Index_j^k|$, with $k = PDI$, IDV, MAS, UAI, LTO, IVR.[5]

4. Results

We focus on the first five legislatures. Data pertaining to roll calls, national parties and European political groups have been obtained from http://personal.lse.ac.uk/hix/HixNouryRolandEPdata.HTM. From the agreement matrix based on data on roll call votes we obtained estimates of legislators' bliss points on the two dimensions of the policy space, by implementation of the DW-NOMINATE. Similarly to Hix, Noury and Roland (2006), we construct our dependent variables as the averages of the positions of legislators belonging to the same national party. There-

[5] Cultural indexes IDV, MAS and UAI are available separately for French Belgium and Flemish Belgium, which have been treated as separate countries.

fore, our analysis focuses on national party characteristics rather than on individual legislators' features. In order to take advantage of a larger dataset, we stack data for five legislatures. We thus obtain two sets of dependent variables, denoted by y_d, where $d = 1, 2$ indicates dimension. We drop from the sample observations for which the full set of variables, required to carry out the analysis described below, is not available. Thus our dataset consists of a total of 347 data points.[6] In order to avoid repetitions, we refer to Hix, Noury and Roland (2006) for an exhaustive descriptive analysis of roll call votes data and of legislators' respective positions in the political space.

The main scope of this empirical analysis is to investigate whether and how correlations among legislators matter in explaining their relative positioning in the political spectrum. Hence, we begin our analysis by looking for the presence of spatial correlation across y_d along the two dimensions. We perform a Moran I test (Moran, 1950), which is designed to detect spatial clustering of data according to some measures of proximity as given by the choice of W. The null hypothesis of the Moran I test, which in its scaled version has been shown to be equivalent to a Lagrange Multiplier test (Burridge, 1980), is the lack of spatial correlation across data. Thus, if the value of the Moran I statistic exceeds the corresponding χ^2 critical value we reject the null hypothesis and conclude that the data are affected by spatial correlation. The value of Moran I statistics are reported in Table 1. Results supports our conjecture of a strong spatial network effect, particularly on the second dimension.

Table 1: Moran I test based on raw data.

	First Dimension	**Second Dimension**
Km	3.64*	120.68***
Lang.	1.22	48.56***
Inst.	6.95***	102.38***
PDI	3.73*	81.98***
IDV	1.69	96.32***
MAS	0.91	53.99***
UAI	1.33	57.31***
LTO	6.66***	98.00***
IVR	1.30	90.49***

Notes. * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

[6] Specifically, we have 43, 55, 62, 81 and 106 observations in the first, second, third, fourth and fifth parliament, respectively.

We next replicate the analysis in Hix, Noury and Roland (2006) to test whether the aforementioned spatial correlation can be fully explained by an appropriate set of regressors or if a spatial model such as (1) is needed. Let LR and $EUint$ be indexes of left-right political orientation and EU integration propensity, respectively. Let D be a set of dummy variables containing country-specific and European political group-specific controls, as well as dummy variables to indicate whether the national party was in power during each legislature (taking value one if the national party was in power for the majority of the legislature and zero otherwise), and whether it had a European Commissioner during such period of time (taking value one if it had a Commissioner for the whole period, 0.5 if it had a Commissioner for at least half of the period, and zero otherwise). LR and $EUint$ have been obtained from expert judgement data in Marks and Steenbergen (2004), while the dummy variables have been obtained from information contained in the European Parliament and European Commission websites, as well as in http://personal.lse.ac.uk/hix/HixNouryRolandEPdata.HTM.

We estimate the parameters of the following regression

$$y_d = \beta_0 + \beta_1 LR + \beta_2 EUint + \gamma D + \epsilon \tag{3}$$

and perform a Moran I test on the obtained Ordinary Least Squares residuals. The value of Moran I statistics reported in Table 2 indicate that residuals from regressions along both dimensions display severe spatial correlation for almost all the choices of proximity measures. This means that the exogenous regressors are not able to account for spatial patterns in the dependent variables.

Table 2: Moran I test based on the specification of Hix et al (2006).

	First Dimension	**Second Dimension**
Km	11.45***	24.54***
Lang.	2.42	25.78***
Inst.	7.14***	13.01***
PDI	6.23***	21.12***
IDV	4.80**	34.97***
MAS	14.41***	4.71**
UAI	10.25***	19.20***
LTO	2.90*	37.98***
IVR	7.48***	25.08***

Notes. * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

Table 2 motivates the inclusion of explicit spatial components into the regression equations. Our main specification of (2) is therefore

$$y_d = \lambda W y_d + \beta_1 LR + \beta_2 EUint + \beta_3 W * LR + \beta_4 W * EUint + \gamma D + \epsilon, \qquad (4)$$

where variables are defined as in (3). As previously mentioned we pool data pertaining to different legislatures, but W is constructed so that spatial correlation across observations only affects units within the same legislature. Thus, all our choices of W have a block diagonal structure where each block reflects interactions of agents within each legislature. A preliminary analysis reveals that the two exogenous effects $W * LR$ and $W * EUint$ are highly correlated, hence we only include one of the two exogenous components in order to avoid inflated standard errors. Specifically, the terms $W * LR$ and $W * EUint$ are separately included in the analysis of the first and second dimensions, respectively.[7] Results are reported in Tables 3 and 4.

The baseline specification in (4) does not include any dummy variable to control for time trends across European countries. We expect global political trends in Europe to generate ex-ante correlations across legislators along unobservable characteristics. Such trends may induce correlations in legislators' positions that are independent from their interactions within the European Parliament itself. We therefore consider an alternative specification as

$$y_d = \lambda W y_d + \beta_1 LR + \beta_2 EUint + \beta_3 W * LR + \beta_4 W * EUint + \gamma D + \delta P + \epsilon, \qquad (5)$$

where P is a set of dummy variables that controls for the legislature.[8] Results are reported in Tables 5-8. Specification (5) allows us to isolate the effects of spatial correlations within each legislature from the effects of the time-varying composition of the Parliament. Hence, it is the most appropriate model to investigate how possible interactions of legislators along spatial characteristics affect their positioning in the European policy space.

Estimates reported in Tables 3-6 have been obtained by a maximum likelihood principle. For each table in the following sections we also report the value of the Lagrange Multiplier statistics (LM henceforth) to test the null hypothesis that the residuals computed for each specification are free from spatial correlation. If LM is not significant we can then conclude that all sources of spatial interactions have been controlled for and are generated by observable channels, either endogenously through the lagged dependent variable or exogenously via the lagged regressors $W * LR$ and/or $W * EUint$.

4.1. Geopolitical proximities

We begin the analysis of our results by considering the correlations induced by our geopolitical proximity matrices, i.e. W^G (geographical), W^L (linguistic) and W^I (in-

[7] This choice is motivated by the fact that LR represents the main explanatory variable along the first dimension, while $EUint$ is only relevant for the second dimension.

[8] The first European Parliament is considered as reference group.

stitutional) using the specification given in (4).

The first dimension, as in Hix, Noury and Roland (2006), is essentially interpreted as the ideological position on the left-right scale. Results reported in Table 3 display an endogenous spatial effect, measured by the estimate of λ, when W is based on geographical and institutional distances. The second dimension, instead, is explained by the positioning on both the left-right scale and the EU integration scale (consistently with Hix, Noury and Roland, 2006). The estimate of λ reported in Table 4 is statistically significant for all choices of W, revealing the presence of a strong endogenous network effect on the second dimension. We outline that results displayed in Tables 3 and 4 are consistent with those of Hix, Noury and Roland (2006), as the coefficients of their main regressors (i.e., LR and $EUint$) have the same sign and level of significance. This holds regardless of which choice of W we adopt.

As previously discussed, specification (4) does not control for the legislature. However, every election changes the identity of legislators. If unobserved time trends lead spatially close countries to elect candidates who are similar to each other, this will result in spuriously inflated spatial parameter estimates. Indeed, the magnitude of the estimates of the spatial parameters erroneously account for both these time trends and the within-parliament spatial correlation generated by the various notions of proximity. We therefore perform a similar analysis controlling for legislatures, in order to isolate the effects of spatial correlations for a given set of legislators, as illustrated by model (5).

Tables 5 and 6 display estimates of the parameters in model (5). Once we take into account possible political trends by controlling for the legislature, the spatial component along the first dimension disappears, as the estimate of λ is never significant. However, the estimate of β_3 is significant, revealing an exogenous network effect driven by LR. Thus, the comparison between Table 3 and Table 5 suggest that the endogenous spatial effects of Table 3 are caused by the spatial correlations that are originated by unobservable political trends that may lead to the election of similar (or dissimilar) candidates rather than by genuine spatial inter-connections across legislators.

Figures in Table 6 show that the institutional weight matrix W^I is the only one that induces a strongly significant endogenous effect along the second dimension. The endogenous effect captures a channel through which the spatial correlations affect legislators' positioning that is not related to the other exogenous variables we included. The coefficient λ is significant and negative. The negative estimate of λ when contiguity is embedded in the matrix W^I is consistent with the existence of a convergence process, as legislators with a different institutional background tend to move their bliss points closer to each other after gaining increasing knowledge of their neighbours' preferences. We interpret this endogenous effect as resulting from a reciprocal influence that induces legislators to modify their preferences over the length of the legislature. The building block of our empirical strategy is the estimate of a single legislators' bliss point per

Table 3: Geopolitical proximities: first dimension, main specification (no legislature dummies).

First Dimension	Km (1)	Lang. (2)	Inst. (3)
λ	0.3057	0.0933	0.3221
	(2.14)**	(0.75)	(2.40)**
LR	1.0853	1.0856	1.0777
	(14.64)***	(14.60)***	(14.51)***
EUint	0.0095	0.0098	0.0090
	(1.04)	(1.07)	(0.99)
$W * LR$	-0.1361	-0.1439	0.1672
	(-0.42)	(-1.37)	(1.06)
EP	No	No	No
EPG	Yes	Yes	Yes
Const.	-0.0800	-0.0366	-0.2206
	(-0.43)	(-0.25)	(-1.45)
LM	5.79**	1.55	2.98*
N	347	347	347

Notes. t-statistics in parentheses. * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

legislature, where each legislature spans a period of five years. In order to fully confirm our dynamic interpretation we would need estimates of legislators' positions on a finer temporal scale within each legislature. In Section 6 we discuss more extensively how a different approach is needed to test whether time-varying preferences are reinforcing (or even driving, in some cases) positive or negative cross-correlation effects.

4.2. Cultural proximities

We replicate the same analysis with a focus on the cultural proximities based on the work by Hofstede, Hofstede and Minkov (2010). We discussed in Section 4.1 how specification 5 is the most appropriate to analyse the effects of spatial correlations within the length of the legislature, isolating them from the political trends across Europe which may as well have a spatial component. For this reason, we only report results of the analysis performed with specification (5) for the cultural proximity matrices.

Table 7 shows results for the first dimension of the policy space. As for the geopolitical proximities, the first dimension is still explained mostly by the ideological positioning

Table 4: Geopolitical proximities: second dimension, main specification (no legislature dummies).

Second Dimension	Km (1)	Lang. (2)	Inst. (3)
λ	0.5115	0.3836	0.4018
	(4.17)***	(3.46)***	(2.71)***
LR	-0.6853	-0.6673	-0.6919
	(-5.08)***	(-4.92)***	(-5.03)***
EUint	0.0300	0.0282	0.0294
	(1.81)*	(1.70)*	(1.75)*
$W * EUint$	0.1017	-0.0255	-0.0148
	(1.58)	(-1.24)	(-0.56)
EP	No	No	No
EPG	Yes	Yes	Yes
Const.	-0.2797	0.3574	0.3095
	(-0.76)	(1.30)	(1.16)
LM	0.44	1.40	0.04
N	347	347	347

Notes. t-statistics in parentheses. * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

of the national party on the left-right scale. We observe that the estimate of β_3 is significant in columns (1) and (5), revealing an exogenous network effect depending on LR. On the other hand, the estimate of λ is significant in column (2), suggesting that, unlike in the results displayed in Table 5, there is significant spatial correlation due to endogenous channels along the first dimension when contiguity is defined according to the IDV index. The IDV index measures whether individuals of a country are expected to take care only of themselves and their immediate families, or whether there is in-group loyalty. This index is correlated to attitudes towards several policies such as healthcare and, more generally, welfare, that are typically related to the left-right orientation. Kallio and Niemelä (2014), for example, show that the left oriented individuals are less prone to display individualistic attitudes in the attribution of poverty. Hence, we interpret the significance of the IDV index as an underlying spatial correlation among countries in unobservable characteristics that are related once more to the positioning on the LR scale.

Table 8 reports the results for the second dimension, which is partially explained

Table 5: Geopolitical proximities: first dimension, main specification, legislature dummies.

First Dimension	Km (1)	Lang. (2)	Inst. (3)
λ	-0.2723	-0.0339	0.0850
	(-1.08)	(-0.26)	(0.48)
LR	1.0567	1.0591	1.0499
	(14.65)***	(14.59)***	(14.43)***
EUint	0.0125	0.0124	0.0113
	(1.40)	(1.39)	(1.27)
$W * LR$	-1.0252	-0.1782	0.3767
	(-2.62)***	(-1.72)*	(2.19)**
EP	Yes	Yes	Yes
EPG	Yes	Yes	Yes
Const.	0.1449	-0.1244	-0.5087
	(0.73)	(-0.88)	(.3.16)***
LM	1.18	0.96	0.17
N	347	347	347

Notes. t-statistics in parentheses. * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

by LR and $EUint$, as in section 4.1. Column (3) shows a significant endogenous spatial effect of W^{MAS}, which is the proximity matrix based on the Masculinity index. The Masculinity index is associated with the importance attached to goal items such as earnings and advancement to higher level jobs (as opposed to a good working relationship with superiors and cooperation with co-workers). The index was labelled Masculinity index because it describes the only cultural dimension in which male and female respondents scored consistently differently (Hofstede, Hofstede, Minkov, 2010, p.139). The Masculinity Index, moreover, is correlated with political outcomes such as female participation in politics. Therefore, in Section 4.3, we further investigate whether the endogenous effect of W^{MAS} may depend on heterogeneous gender composition of national parties. Additionally, column (3), (5) and (6) show that there is an exogenous spatial effect of W^{MAS}, W^{LTO}, and W^{IVR}.

Table 6: Geopolitical proximities: second dimension, main specification, legislature dummies

Second Dimension	Km (1)	Lang. (2)	Inst. (3)
λ	-0.3946	0.0824	-0.6953
	(-1.57)	(0.60)	(-3.26)***
LR	-0.7259	-0.7171	-0.7328
	(-5.47)***	(-5.38)***	(-5.57)***
EUint	0.0352	0.0331	0.0338
	(2.15)**	(2.02)**	(2.10)**
$W * EUint$	0.0610	-0.0033	0.0372
	(0.89)	(-0.16)	(1.24)
EP	Yes	Yes	Yes
EPG	Yes	Yes	Yes
Const.	-0.2490	0.0699	-0.1490
	(-0.66)	(0.25)	(-0.49)
LM	1.45	0.15	0.00
N	347	347	347

Notes. t-statistics in parentheses. * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

4.3. Effects of gender composition

Results displayed in Table 8 suggest that gender composition of national parties may play a role in explaining the determinants of the European policy space. We therefore investigate whether including gender composition as a regressor is enough to account for the strong significance of the estimate of λ when contiguity is defined according to W^{MAS}. We define a new independent variable G as the percentage of female legislators in the national party in the chosen legislature. Data about gender of legislators have been deduced from the European Parliament website. We therefore introduce an extended specification as

$$y_d = \lambda W y_d + \beta_1 LR + \beta_2 EUint + \beta_3 W * LR + \beta_4 W * EUint + \beta_5 G + \beta_6 W * G + \gamma D + \delta P + \epsilon. \quad (6)$$

Once again we include either $W * LR$ or $W * EUint$ depending on which dimension we are analysing in order to mitigate the issue of multicollinear regressors. Correlations between $(W * LR)/(W * EUint)$ and $W * G$ instead do not pose any issue of collinearity. Results for geopolitical proximities, W^{MAS} and W^{IDV} are reported in Tables 9 and

Table 7: Cultural proximities: first dimension, main specification, legislature dummies.

First Dimension	PDI (1)	IDV (2)	MAS (3)	UAI (4)	LTO (5)	IVR (6)
λ	-0.0056	-0.4811	0.0421	0.2023	-0.0692	-0.0642
	(-0.04)	(-2.10)**	(0.30)	(1.39)	(-0.42)	(-0.38)
LR	1.0563	1.0607	1.0667	1.0627	1.0740	1.0649
	(14.56)***	(14.63)***	(14.68)***	(14.65)***	(14.94)***	(14.61)***
EUint	0.0131	0.0115	0.0115	0.0122	0.0114	0.0122
	(1.47)	(1.29)	(1.28)	(1.37)	(1.29)	(1.37)
$W * LR$	-0.5028	0.0913	0.2264	0.2036	-0.5051	-0.1342
	(-2.08)**	(0.33)	(1.06)	(1.20)	(-3.02)***	(-0.57)
EP	Yes	Yes	Yes	Yes	Yes	Yes
EPG	Yes	Yes	Yes	Yes	Yes	Yes
Const.	-0.1512	-0.3408	-0.3001	-0.3799	0.1504	-0.2137
	(-1.20)	(-2.34)**	(-2.54)***	(-2.47)**	(0.85)	(-1.42)
LM	0.64	0.71	0.26	0.30	0.85	0.02
N	347	347	347	347	347	347

Notes. t-statistics in parentheses. * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

10.[9]

Compared with the results displayed in Tables 5 and 7, the inclusion of G does not affect either the interpretation of the first dimension as left-right orientation or the significance of the endogenous effect obtained with W^{IDV}. The estimate of β_6 is significant only when contiguity is defined according to W^{MAS}, revealing a direct effect of G on neighbouring countries when neighbours are defined in terms of MAS index. On the other hand, gender composition of the national party affects significantly the positioning on the second dimension, as β_5 is always significantly positive. Moreover, the spatial endogenous effects of the institutional proximity remains significant, and the endogenous effect of W^{MAS} even more so revealing that positions of legislators in the European Parliament are not only affected by gender composition, but also by an endogenous network channel based on W^{MAS}.

[9]Results for the additional cultural proximities are available on request from the Authors.

Table 8: Cultural proximities: second dimension, main specification, legislature dummies.

Second Dimension	PDI (1)	IDV (2)	MAS (3)	UAI (4)	LTO (5)	IVR (6)
λ	-0.2153	0.1659	-0.3889	-0.0414	0.2268	0.0480
	(-1.25)	(0.90)	(-2.46)**	(-0.26)	(1.46)	(0.28)
LR	-0.7302	-0.7146	-0.7176	-0.7253	-0.7133	-0.7086
	(-5.48)***	(-5.34)***	(-5.47)***	(-5.43)***	(-5.40)***	(-5.36)***
EUint	0.0343	0.0334	0.0330	0.0332	0.0331	0.0345
	(2.09)**	(2.03)**	(2.04)**	(2.03)**	(2.04)**	(2.13)**
$W * EUint$	0.0082	-0.0124	0.0774	0.0242	-0.0658	0.0980
	(0.20)	(-0.27)	(1.95)*	(0.88)	(-2.34)**	(2.54)***
EP	Yes	Yes	Yes	Yes	Yes	Yes
EPG	Yes	Yes	Yes	Yes	Yes	Yes
Const.	-0.0017	0.1178	-0.2011	-0.1035	0.6383	-0.3790
	(-0.05)	(0.42)	(-0.90)	(-0.37)	(2.02)**	(-1.40)
LM	0.55	2.75*	0.56	0.00	0.30	0.71
N	347	347	347	347	347	347

Notes. t-statistics in parentheses. * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

5. Robustness checks

5.1. Spectral norm normalisation

As mentioned in Section 2, the weight matrix W is chosen ex-ante by the practitioner and scaled to guarantee reliability of standard estimation methods.[10] The row normalisation we adopt is widely used in the spatial econometrics literature and allows a weighted average interpretation of the spatial autoregressive terms, i.e. each unit is possibly related to a weighted average of neighbouring units. In our case this means that every national party is overall equally influenced by all the other national parties in the European Parliament, and the heterogeneity rests only in the relative weights with which different parties are influenced by their neighbours. However, one can imagine that more peripheral parties are influenced less (or more) by their neighbours than those with a central position in the economic space considered. Hence, we run a robust-

[10] Typical technical issues are for instance existence of the likelihood function and parameters' identification.

Table 9: First dimension, gender effects, legislature dummies.

First Dimension	Km (1)	Lang. (2)	Inst. (3)	IDV (4)	MAS (5)
λ	-0.2742	-0.0425	0.0922	-0.4939	0.0345
	(-1.08)	(-0.33)	(0.50)	(-2.10)**	(0.25)
LR	1.0542	1.0582	1.0462	1.0569	1.0429
	(14.61)***	(14.55)***	(14.35)***	(14.54)***	(14.29)***
EUint	0.0130	0.0123	0.0115	0.0117	0.0106
	(1.47)	(1.37)	(1.29)	(1.31)	(1.18)
$W * LR$	-0.9589	-0.2254	0.3966	0.1164	-0.0316
	(-2.38)**	(-1.47)	(-0.13)	(0.39)	(0.05)
G	-0.0217	-0.0198	-0.0305	-0.0239	-0.0222
	(-0.57)	(-0.52)	(-0.79)	(-0.63)	(-0.58)
$W * G$	-0.2084	0.0821	-0.0118	-0.0582	0.4916
	(-0.57)	(0.44)	(-0.03)	(-0.20)	(2.17)**
EP	Yes	Yes	Yes	Yes	Yes
EPG	Yes	Yes	Yes	Yes	Yes
Const.	0.1538	-0.1320	-0.5143	-0.3426	-0.2256
	(0.78)	(-0.93)	(-3.19)***	(-2.35)**	(-1.84)*
LM	1.05	1.28	0.18	0.69	0.13
N	347	347	347	347	347

Notes. t-statistics in parentheses. * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

Table 10: Second dimension, gender effects, legislature dummies.

Second Dimension	Km (1)	Lang. (2)	Inst. (3)	IDV (4)	MAS (5)
λ	-0.2852	0.0666	-0.6349	-02941	-0.4022
	(-1.07)	(0.47)	(-2.81)***	(-1.27)	(-2.53)**
LR	-0.7077	-0.6977	-0.7049	-0.6793	-0.6973
	(-5.38)***	(-5.27)***	(-5.39)***	(-5.18)***	(-5.30)***
EUint	0.0345	0.0314	0.0327	0.0321	0.0318
	(2.13)**	(1.93)*	(2.05)**	(2.00)**	(1.98)**
$W * EUint$	0.0649	-0.0080	0.0568	-0.7382	0.0763
	(0.94)	(-0.31)	(1.29)	(-1.37)	(1.84)*
G	0.1705	0.1814	0.1652	0.1900	0.1811
	(2.45)**	(2.61)***	(2.39)**	(2.76)***	(2.65)***
$W * G$	-0.5280	0.0798	-0.5145	1.8138	-0.0521
	(-0.77)	(0.25)	(-0.85)	(3.14)***	(-0.14)
EP	Yes	Yes	Yes	Yes	Yes
EPG	Yes	Yes	Yes	Yes	Yes
Const.	-0.1852	0.0596	-0.1154	0.1046	-0.2206
	(-0.48)	(0.21)	(-0.38)	(0.40)	(-0.97)
LM	1.05	0.08	0.01	2.17	0.56
N	347	347	347	347	347

Notes. t-statistics in parentheses. * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

ness check of our last specification given in (6) where, for every weight matrix, we scale each element of W by its spectral norm, rather performing a row-normalisation.[11] The spectral norm normalisation has the advantage of preserving in full the heterogeneity across different rows, as all elements of W are scaled by the same factor. Results in Tables 11 and 12 (in the Appendix) show that our main results still hold, with the exceptions of the loss of significance of the estimate of λ along the second dimension when contiguity is defined according to MAS and its weak significance (as opposed to no significance) when W^{IDV} is chosen. The significance of W^{IDV} may be related once more to the fact that this proximity matrix appears to capture correlation on characteristics related to the LR index, which is relevant also on the second dimension.

5.2. Spatial lagged independent variables

Spatial autoregressions such as (1) and (2) have been criticized by applied researchers (e.g., Gibbons and Overman, 2012) because of their peculiar functional form and the ex-ante choice of the spatial weights. More specifically, in case the spatial weights are not correctly specified, identification issues might arise, leading then to spurious estimates and misleading inference. In order to assess the reliability of our results, we focus on the alternative simpler model

$$y_d = \beta_1 LR + \beta_2 EUint + \beta_3 W * LR + \beta_4 W * EUint + \beta_5 G + \beta_6 W * G + \gamma D + \delta P + \epsilon. \quad (7)$$

generally known as the spatial lag of X (SLX) model. All the variables are defined as in Section 4. As previously discussed, the terms $W * LR$ and $W * EUint$ have been included separately when considering the first and second dimension, respectively. Comparison of Tables 5-6 and 15-16 (reported in the Appendix) reveals that the estimates of β_1-β_6 retain almost always the same level of significance and roughly the same magnitude. This supports the robustness of our main results, and suggests that the estimates of the spatial coefficient λ provide meaningful insights in analyzing the position of legislators along the second dimension of the political space without affecting the direct and exogenous effects.

6. Conclusions and future outlooks

The main scope of this paper is to evaluate whether heterogeneity and spatial correlations across legislators do affect their positioning in the European policy space, and henceforth the interpretation of the dimensions of the policy space itself. We extend the analysis of Hix, Noury and Roland (2006) by incorporating spatial econometrics techniques into the standard regression methods to investigate the determinants of legislators' behaviour. More specifically, we investigate the channels of spatial correlation

[11] The spectral norm is defined as the square root of the maximum eigenvalue of $W'W$, where prime denotes transposition.

across MEPs, where distances are intended as measures of political, geographical and cultural proximities.

Our empirical analysis confirms that the first dimension of the political space is mainly explained by the left-right political orientation, consistently with previous findings in the literature. However, it is also clear from our results that omitting spatial components in the analysis leads to incomplete results, as political orientation of neighbouring countries affects legislators' positions for several measures of proximities (i.e., W^G, W^I, W^L, W^{PDI} and W^{LTO}, as clearly shown by Tables 5 and 7). Moreover, we identify a robust endogenous channel of spatial correlation across legislators along the first dimension of the political space when the weight matrix is based on the IDV index. The IDV index captures characteristics of the country that are also related to left-right ideological orientation, such attitudes towards healthcare provisions, or more generally towards welfare. We therefore conclude that the first dimension is indeed explained by the ideological positioning on the LR scale, and by those spatial effects, either exogenous or endogenous, that are related to LR.

Our analysis of the second dimension, instead, reveals that both LR and $EUint$ play a role, in line with previous literature. Also, W^I, which is based on an institutional distance measure, induces a strong correlation across legislators through an endogenous channel, suggesting the presence of a convergence process that requires further investigation. In addition, we find a strongly significant spatial effect of the endogenous type when the weight matrix is based on the MAS index. However, the latter is not robust to the change of the normalisation factor of the weight matrix. Results in Table 8 also reveal a significant direct effect of European integration of neighbouring countries on legislators positions when distances are defined according to W^{MAS}, W^{LTO} and W^{IVR}.

A further novel contribution of this paper is the inclusion of gender composition as one of the potential determinants of legislators' positions along the two dimensions. We find that the percentage of women in national parties plays a role to explain the second dimension, without affecting the significance and magnitude of the aforementioned analysis. Instead, our gender variable plays no role in explaining the first dimension of the political spectrum, as somewhat expected.

Findings of this paper open up several lines of research that await to be tackled. First, as discussed in Section 4, the strongly significant negative sign of the endogenous spatial component in our regressions suggests a convergence process of legislators' ideology within each legislature. If this holds true, we should find a stronger spatial effect when the legislature approaches its end, compared to its beginning. However, as in the current setting we obtain one estimate of each legislator's bliss point per legislature and such estimates are then aggregated at national party level, we cannot formally test whether this is the case. We plan to extend our analysis to a finer scale by considering observations pertaining to individual legislators rather than working at national party level. We can thus exploit a larger number of observations and explore further the transmission mechanisms of the effects we have found. We also aim to estimate several legislators positions during the same legislature, so to be able to understand whether convergence is the true mechanism that induces a significant spatial correlations across legislators through endogenous channels.

A second crucial point we plan to address in the near future is methodological and involves the starting point of our analysis, i.e. the NOMINATE procedure and its variants (such as DW-NOMINATE). The NOMINATE technique, as it is currently known and used in the literature, estimates parameters of a utility function for legislator i when he/she votes *yes* in roll call j. The latter is defined as (e.g., Poole, 2005)

$$U_{i,j,yes} = u_{i,j,yes} + \nu_{i,j,yes}, \tag{8}$$

where $u_{i,j,yes}$ is the deterministic component and $\nu_{i,j,yes}$ is a random shock. Similarly to Poole (2005), the random components $\nu_{i,j,yes}$ (and correspondingly $\nu_{i,j,no}$) are independent (across i and j) and identically distributed random draws from a normal distribution. However, this choice overlooks potential *a priori* spatial correlation and heterogeneity of unknown form across legislators, leading to inaccurate estimates. Hence, a crucial methodological point that needs to be addressed in order to improve results of this paper and, more generally, to open up new frontiers in the analysis of the political space, is the extension of NOMINATE to estimate parameters of a random utility function where random shocks are possibly heterogeneous and spatially correlated. Since the findings of this paper clearly identify several different geopolitical and cultural distances that might be relevant to explain legislators' positions, we anticipate that the extension to NOMINATE should allow for an *ex-ante* heterogeneity and spatial correlation of unknown form in order to refine the estimates of legislators' ideal points in the political space. We can *ex-post* perform an analysis similar to that carried out in this paper to investigate explicitly different channels of correlation by means of particular notions of distances.

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A. Additional tables

Table 11: First dimension with spectral norm normalisation.

First Dimension	Km (1)	Lang. (2)	Inst. (3)	IDV (4)	MAS (5)
λ	-0.2249	-0.2096	0.3035	-1.0290	-0.0149
	(-0.68)	(-1.00)	(1.18)	(-2.84)***	(-0.06)
LR	1.0674	1.0629	1.0560	1.0601	1.0320
	(14.69)***	(14.59)***	(14.39)***	(14.72)***	(14.25)***
EUint	0.0123	0.0119	0.0123	0.0121	0.0110
	(1.38)	(1.33)	(1.37)	(1.37)	(1.25)
$W * LR$	0.9858	-0.2803	0.2702	0.9371	0.3066
	(1.36)	(-0.82)	(0.58)	(1.53)	(0.99)
G	-0.0306	-0.0228	-0.0240	-0.0312	-0.0325
	(-0.79)	(-0.60)	(-0.62)	(-0.82)	(-0.86)
$W * G$	-0.5561	0.3785	-0.0474	-0.9509	0.6157
	(-0.95)	(1.05)	(-0.09)	(-1.56)	(1.79)
EPG	Yes	Yes	Yes	Yes	Yes
EP	Yes	Yes	Yes	Yes	Yes
Const.	-0.3858	-0.2812	-0.3169	-0.3104	-0.1819
	(-2.77)***	(-2.34)**	(-2.32)**	(-2.72)***	(-1.59)
LM	0.72	0.68	0.05	0.27	0.03
N	347	347	347	347	347

Notes. t-statistics in parentheses. * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

Table 12: Second dimension with spectral norm normalisation.

Second Dimension	Km (1)	Lang. (2)	Inst. (3)	IDV (4)	MAS (5)
λ	-0.4155	0.0158	-0.7888	-0.7714	-0.3865
	(-1.11)	(0.06)	(-2.21)**	(-1.95)*	(-1.34)
LR	-0.7079	-0.6934	-0.6671	-0.6735	-0.7029
	(-5.40)***	(-5.25)***	(-5.08)***	(-5.16)***	(-5.27)***
EUint	0.0347	0.0308	0.0312	0.0324	0.0314
	(2.15)**	(1.90)*	(1.95)*	(2.02)**	(1.94)*
$W * EUint$	0.2296	0.0034	-0.0607	0.0340	0.0807
	(1.74)*	(0.06)	(-0.76)	(0.29)	(1.60)
G	0.1509	0.1775	0.1740	0.1781	0.1757
	(2.15)**	(2.56)***	(2.52)**	(2.60)***	(2.54)***
$W * G$	-1.7913	0.3840	0.0543	1.5504	-0.0506
	(-1.74)*	(0.71)	(0.06)	(1.53)	(-0.09)
EP	Yes	Yes	Yes	Yes	Yes
EPG	Yes	Yes	Yes	Yes	Yes
Const.	-0.2668	-0.0454	0.4152	0.0437	0.0146
	(-0.92)	(-0.19)	(1.60)	(0.20)	(0.07)
LM	0.87	0.79	0.04	1.70	1.25
N	347	347	347	347	347

Notes. t-statistics in parentheses. * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

Table 13: First dimension, SLX model

First Dimension	Km (1)	Lang. (2)	Inst. (3)	IDV (4)	MAS (5)
LR	1.0561	1.0590	1.0464	1.0621	1.0435
	(14.60)***	(14.56)***	(14.34)***	(14.49)***	(14.31)***
EUint	0.0127	0.0123	0.0116	0.0121	0.0105
	(1.43)	(1.37)	(1.31)	(1.35)	(1.18)
$W * LR$	-0.8922	-0.2250	0.4015	0.1509	-0.0359
	(-2.21)**	(-1.47)	(1.58)	(0.51)	(-0.15)
G	-0.0211	-0.0205	-0.0301	-0.0223	-0.0213
	(-0.55)	(-0.51)	(-0.78)	(-0.58)	(-0.56)
$W * G$	-0.2110	0.0719	-0.0374	0.0427	0.4944
	(-0.58)	(0.38)	(-0.11)	(0.15)	(2.19)**
EPG	Yes	Yes	Yes	Yes	Yes
EP	Yes	Yes	Yes	Yes	Yes
Const.	0.1443	-0.1232	-0.5113	-0.3177	-0.2264
	(0.73)	(-0.87)	(-3.23)***	(-2.17)**	(-1.85)*
N	347	347	347	347	347

Notes. t-statistics in parentheses. * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

Table 14: Second dimension, SLX model

Second Dimension	Km (1)	Lang. (2)	Inst. (3)	IDV (4)	MAS (5)
LR	-0.7054	-0.7010	-0.7040	-0.6761	-0.6988
	(-5.34)***	(-5.30)***	(-5.30)***	(-5.14)***	(-5.25)***
EUint	0.0337	0.0311	0.0317	0.0319	0.0309
	(2.07)**	(1.91)*	(1.95)*	(1.98)**	(1.90)*
$W * EUint$	0.3921	-0.1965	0.6157	-0.6542	0.4741
	(0.53)	(-0.71)	(1.33)	(-1.23)	(1.07)
G	0.1690	0.1837	0.1641	0.1827	0.1763
	(2.42)**	(2.65)***	(2.34)**	(2.67)***	(2.55)***
$W * G$	-0.8232	0.2788	-0.8083	1.4295	-0.1434
	(-1.24)	(0.81)	(-1.33)	(2.81)***	(-0.35)
EP	Yes	Yes	Yes	Yes	Yes
EPG	Yes	Yes	Yes	Yes	Yes
Const.	0.0238	0.0487	-0.0605	0.1209	-0.0379
	(0.07)	(0.19)	(-0.21)	(0.46)	(-0.17)
N	347	347	347	347	347

Notes. t-statistics in parentheses. * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$