# Supporting Information Online

# Networks and Social Influence in European Legislative Politics

Thomas Malang[a] Laurence Brandenberger[b,c] Philip Leifeld[d]

[a]University of Konstanz, Department of Politics & Public Administration

[b]Eawag (Swiss Federal Institute of Aquatic Science and Technology)

[c]University of Bern, Institute of Political Science

[d]University of Glasgow, School of Social and Political Sciences

# 1 Specification of control variables

We control for two alternative homophily patterns.

First, we expect variables related to the political system to shape the preferences of a parliament towards EU policy. Political institutions matter because similar institutions presumably lead to similar policy preferences and similar kinds of transaction costs (Bennett 1991). In parliamentary systems, the majority party and the government party are usually identical, but this congruence is not necessarily given in presidential systems. We expect chambers in presidential systems to enjoy a greater degree of freedom and to place a greater emphasis on scrutinizing the government. Therefore, chambers from the same type of political system should act in similar ways (selection), and there may even be coordination or imitation among presidential systems or among parliamentary systems at a higher rate than across these two types (social influence). We test political system homophily (*Institutional homophily*) using a node-match term (Equation 3) based on the "political system" variable from Armingeon's Comparative Political Data Set III (Armingeon et al. 2016).

Second, Bailer, Mattila and Schneider (2015) argue that economic factors determine deviant voting behavior in the Council. Especially rich and competitive member states are more likely to oppose legislation. These member states use voting against a legislative proposal only to signal their discontent to their domestic stakeholders. We assume that on the parliamentary level richer member states are more likely to veto the same legislative proposals (selection) and coordinate their veto behavior (social influence) to circumvent specific proposals. A sender-sender covariate (Equation 4) is used to capture differences in GDP per capita between the vetoing chambers through the model term *Abs. diff. in GDP*.

Furthermore, *Mean import from indirect ties* is a sender–sender covariate (Equation 4) that controls for trade flows between the countries in which co-vetoing chambers are located in order to account for veto behavior due to trade dependence on other countries. This variable is based on the UN Comtrade database as other frequently used trade datasets in international relations do not include recent years.

Main effects (rather than homophily effects) can be at work at two levels: covariates that increase or decrease the general propensity of parliamentary chambers to issue vetoes, and covariates that increase or decrease the popularity of specific legislative proposals as targets of vetoes. Main effects enter the model weighted by time (see Equation 2). There are several additional variables at the chamber level that were used in previous studies (Gattermann and Hefftler 2015; Auel, Rozenberg and Tacea 2015) or that could have a plausible effect on the odds of a chamber to veto a bill:

First, Neuhold and Strelkow (2012) show that second chambers act more frequently than first chambers. *Second chamber* controls whether second chambers or upper houses differ from first chambers or lower houses in their veto activity.

Second, chambers with more capacity should lead to a higher veto activity. *Capacity* controls for the size of a chamber as measured by its number of seats.

Third, the strength of institutional control rights per chamber influences its overall veto activity. *Control* is taken from Winzen (2012) and captures the actual level of control rights of a chamber, which may be an institutional source of variance in vetoing activity. Note that Croatia joined the European Union in July 2013, hence no recent data about *control* are available for Croatia (= 2.56 percent missing data in the *control* variable). We used multiple imputation based on the other nodal attributes to impute these missing values (Buuren and Groothuis-Oudshoorn 2011).

Fourth, parliaments in countries with more Eurosceptic publics are hypothesized to be more active in vetoing EU legislation. *EU opposition* therefore controls for anti-EU attitudes of a ruling party by their mentioning of EU resentments in party manifestos as recorded by the Manifesto Project (Volkens et al. 2016).

Fifth, we control for the population size of a country in order to account for relative influence as more populous member states' actions may possibly carry more weight in the process (*Population (log)*). This variable controls whether larger countries take on a leading role in the veto process.

Sixth, *Constant GDP per capita* controls whether wealthy or poor countries have different activity levels. *Constant GDP per capita* and *Population (log)* come from the World Bank database and are introduced as main effects. These time-varying data were collected from the World Bank database on a yearly basis for all member states. For some missing values in 2014, the values of 2013 were imputed.

Seventh, *Chamber activity* is a main effect for chambers that sums up the weights of all the past vetoes that the focal chamber was involved in. Chamber activity is included in the model as a *sender activity term*,

$$h^t_{\text{sact}}(i,j) = \sum_{e \in E^t} w_t(i_e, j_e)[i_e = i]. \tag{1}$$

This model term tests whether the probability of vetoing a proposal increases if the chamber has vetoed other proposals in the recent past and therefore controls for differential chamber activity.

At the proposal level, we introduce two additional controls into the model: first, it is important to control for the clustering of multiple legislative proposals around the same chambers based on joint proposal characteristics. We capture this kind of *issue specificity* by controlling for clustering between bills proposed by the same Directorate General (DG) of the European Commission around the same chamber. Issue specificity is a node-match term similar to Equation 3, but with the attribute match occurring at the level of proposals, i. e., $i$ and $j$ are reversed. Issue specificity is included in the model as a *target node-match* term,

$$h^t_{\text{tnm}}(i,j) = \sum_{e \in E^t} w_t(i_e, j_e)[j_e \neq j][i_e = i][x_{j_e} = x_j], \tag{2}$$

where $x_j$ indicates the Directorate General of the Commission that proposed law $j$. This term captures the tendency of actors to engage repeatedly in the same issues by vetoing proposals. We cross-checked every proposal with the EU's EUR-Lex database to get information about the Directorate General (DG) in charge of the proposal. This serves as a measure of the respective policy domain of a proposal (attribute value $x$).

Second, we control directly for the *salience* of an issue by introducing a dummy variable for whether the current legislative proposal is related to agriculture, which is traditionally the most redistributive issue in European politics (Kleine 2013). Salience checks whether a law was proposed by the Directorate General (DG) for Agriculture.

Finally, we include main effects corresponding to the homophily terms. For *Party family*, *Entry round*, and *Political system*, main effects are introduced for the different levels in order to account for such things as nationalist sentiments and core EU members.

# 2 Interpretation of results for the control variables

With regard to the control variables, we see that chambers that play according to the same institutional rules in their respective political system seem to cluster together around specific proposals (as indicated by the original $p$ value for *Institutional homophily*). There is no clear result on whether this is due to social influence or shared underlying traits and interests because the first permutation indicates a significant difference of the homophily pattern from a random temporal sequence while the second permutation indicates no difference between the original coefficient and a model with permuted sequences of events given the global distribution of vetoes across time points (for a visual representation of the permutation effects see Figure 2). If one is willing to make the assumption that parliaments need a great deal of preparation time and can only veto relatively close to the deadline, the large $p$-value for institutional homophily in the second permutation is a hint that social selection, rather than social influence, is the triggering mechanism.

Veto diffusion also takes place among countries with different wealth levels. The larger the difference in GDP between a potential vetoing chamber and a chamber that already issued a veto, the more likely it is that the potential chamber issues a veto as well. The effect is strong and withstands even the second permutation. In comparison with the non-significant result of GDP per capita as a main effect, we get a nuanced understanding of the collective veto action. First, richer countries do not veto with a higher likelihood than poorer member states (a finding that is contrary to the voting pattern in the Council where Bailer, Mattila and Schneider 2015 find that richer member states oppose legislation more often). However, if a parliament from a richer member state initiates a veto, chambers from poorer countries will join with a higher likelihood afterwards. There is some room for the interpretation of this result.

Some chamber characteristics like extensive control rights, cameralism (Neuhold and Strelkow 2012), and nationalist party family majorities cause chambers to veto more proposals than other chambers. Veto diffusion also seems to take place among trade partners, but the pattern is not significant at the 95 percent level. Issue specificity can explain additional variation, which means that chambers tend to veto similar proposals, also in a temporally clustered way. Overall, such main effects seem to play a minor role compared to the homophily effects.

To rule out the possibility that the ideological homophily effect could be driven by a single party family, we introduce dummy variables for all party families as a control (along with similar controls for the other hypotheses). Nationalist parties are skeptical about European integration, which causes parliaments led by nationalist parties to veto more proposals by the Commission.

# 3 Rescaling of endogenous network variables

Endogenous network variables are rescaled using a constant. Without rescaling, the coefficient sizes of the network variables are difficult to interpret.

All homophily variables are rescaled by division by the value $c = 0.035$, which represents the increase in the statistic, given an additional event that occurred ten days before the present event. As all homophily variables are rescaled by the same constant, coefficient sizes for these variables can be compared with each other, the largest coefficient indicating the variable with the largest effect on the probability of event occurrence. Actor activity and issue specificity are rescaled by the same constant.

The constant $c$ is calculated on the basis of the weight function in Equation 2:

$$c = exp(-(t - t_e) \cdot (ln(2)/T_{1/2})) \cdot ln(2)/T_{1/2}$$

$$c = exp(-(10) \cdot (ln(2)/10)) \cdot ln(2)/10$$

$$c = 0.0347$$

Absolute difference in GDP is rescaled by a constant that represents an additional event that occurred ten days in the past and with a difference in GDP per capita of $10,000.

Mean import from indirect ties is rescaled by a constant representing an additional past event that occurred ten days in the past and from whom the focal country imports goods worth one million US-$.

# 4 Timing of vetoes

Figure 1 shows the timing of the individual vetoes for 135 proposals. Five proposals were excluded from the diagram and the analysis as they include nine vetoes that were made over 200 days before the deadline. This seems to be an error in the IPEX database.

# 5 Pairwise correlation matrix

Table 1 reports the pairwise correlations between all continuous model terms. It should be noted that correlations may be high between the homophily variables since they include zero values whenever there has been no past veto that relates to the focal event. There may be excessive correlation among these zero values that distort the overall correlation between the variables.

## 6 Summary statistics

Table 2 reports summary statistics for the variables used in the analysis.

## 7 Outlier bill

Legislative proposal "COD/2015/0070" was excluded from the analysis because it represents an outlier. Table 3 demonstrates that this decision is a) justified (third column) and b) not consequential for the substantive results, other than weakening the effect sizes slightly.

The "DIRECTIVE OF THE EUROPEAN PARLIAMENT AND OF THE COUNCIL amending Directive 96/71/EC of The European Parliament and of the Council of 16 December 1996 concerning the posting of workers in the framework of the provision of services" (COD/2015/0070) deals with defining a set of mandatory rules regarding the terms and conditions of employment to be applied to posted workers. It provides that the principle of equal treatment with local workers also covers posted temporary agency workers, thereby aligning the current legislation on temporary agency work.

As the following quantitative analysis shows, the proposal is a clear outlier in our data. Qualitatively, this outlier status can be backed up by two remarkable facts. First, with 14 parliamentary chambers from 11 Member States, the veto count represents 22 out of 56 votes, the highest amount of votes any individual proposal ever received, from the highest number of individual chambers. But what is most striking is the composition of the parliaments. With the exception of the Danish Folketing, all are Central or Eastern European chambers, forming a "regional block." Most observers agree that economic interests are at the heart of the vetoes. Countries with lower wages dislike the idea of equal wages, as they regard it as a threat to their competitive advantage of being able to pay lower wages. Second, as a result, the Commission maintained the initial proposal intact after the enforced re-examination and overruled the concerns expressed by national parliaments for the first time in the history of the EWS.

Table 3 reports the relational event model including a categorical variable denoting the three proposals that resulted in yellow cards. The veto sequence of the outlier proposal clearly does not follow a party homophily pattern. While the interaction terms for the two other yellow cards show no significant deviation from the significant and positive main effect of party homophily, proposal “COD/2015/0070” shows a negative association for the party homophily pattern (the baseline coefficient plus the negative interaction term results in a ideological homophily coefficient of $-0.13$). It is important to note that the full model including the outlier still yields a significant party homophily effect, with a smaller overall effect size.

# 8 Permutation results for the control variables

Figure 2 depicts the results of the two permutations for the three control variables *ideological homophily*, *absolute difference in GDP per capita* and *mean import from indirect ties.*

The first two variabels have significant and positive coefficients in the REM presented in Table 1. Results from the first permutation round indicate that there is a non-randomness to the order of events. However, when combining the results from the second permutation, only absolute difference in GDP per capita prevails as a temporal pattern.

# 9 Assessment of the presented model

Table 4 compares the full REM presented in the article with a REM including only exogenous covariates. The model improves substantually with the inclusion of the network variables as indicated by the traditional R-squared as well as McFadden’s pseudo R-squared. McFadden pseudo R-squared values are generally low and values of $0.2-0.4$ represent an "excellent fit" (McFadden 1978: 307). Therefore, our value of 0.194 indicates a good fit.

Since most (pseudo) R-squared measures are contested in non-OLS models, Tjur (2009) proposes comparing predictive probabilities across different models. We do this by estimating multiple models, starting with a baseline models containing only fixed effects on party family, entry round and political systems and calculating their predictive power. By adding additional terms to the model and calculating each model's predictive power, we can compare the improvement each term makes to the model. As proposed by Tjur (2009), we use the mean probability of event occurrence (i.e. one minus the probability of surviving; $1 - S(t)$) for the null events and the true events to assess predicitive power.

Figure 3 depics improvement steps graphically. The first larger step in the predictive power is caused by the trade variable. Whereas exogenous chamber or country attributes improve the predictive power of the model only marginally, several endogenous network statistcs prove potent additions to the model. Most notably among them are the absolute difference in GDP per capita as well as the ideological homophily variable. This points to the fact that both party line homophily as well as economic differences are important mechanisms by which vetoing activity is affected.

Overall the predictive power of the model increases from 0.18 to 0.31 with the inclusion of endogenous network variables.

# 10 Comparison with a two-mode ERGM

As a robustness check, Table 5 reports the results of a two-mode exponential random graph model (ERGM) with the same data. The ERGM (Cranmer and Desmarais 2011) is a cross-sectional model, which cannot discriminate between selection and influence. To see this more clearly, consider a temporal sequence of five vetoes where the third actor in the sequence not only considers the two prior events in his or her homophily calculation but also the two posterior events. This may lead to a significant and positive homophily effect, but it is clearly not due to imitation of previous actions of others. It is not possible in an ERGM framework to distinguish between prior and posterior events or any sequence information at all. Therefore this ERGM robustness check can indicate for what variables a homophily effect is present, but it does not tell us whether it is a homophily effect due to selection or due to social influence.

This robustness check is still useful because it is a more established method than the relational event models reported in the main part of the article. If the variables that have a significant result in the REM also have a significant effect in the ERGM, this increases our confidence that the results are valid irrespective of the technique being employed. Table 5 indeed reports very similar results as the REM. If there are any deviations between the REM and the ERGM, then these are cases where the ERGM coefficient is more significant, due to the problem described in the previous paragraph.

The ERGM contains an additional model term for "chamber clustering," which introduces a baseline for the homophily effects. This is not necessary in the REM because this is taken care of at the estimation stage in the conditional logit model. The ERGM reports a model specification that is as close as possible to the REM.

The reduced model in the second column removes some of the model terms that are substantially unimportant for the results, which improves the Bayesian Information Criterion (BIC) somewhat.

Figure 4 reports the goodness of fit statistics of the two-mode ERGM. The sufficient statistics capture the endogenous properties of the network in an adequate way. This follows from the fact that the black line (the observed network statistic) and the boxplots (the same statistic for 1,000 simulated networks from the estimated model) are nearly identical.

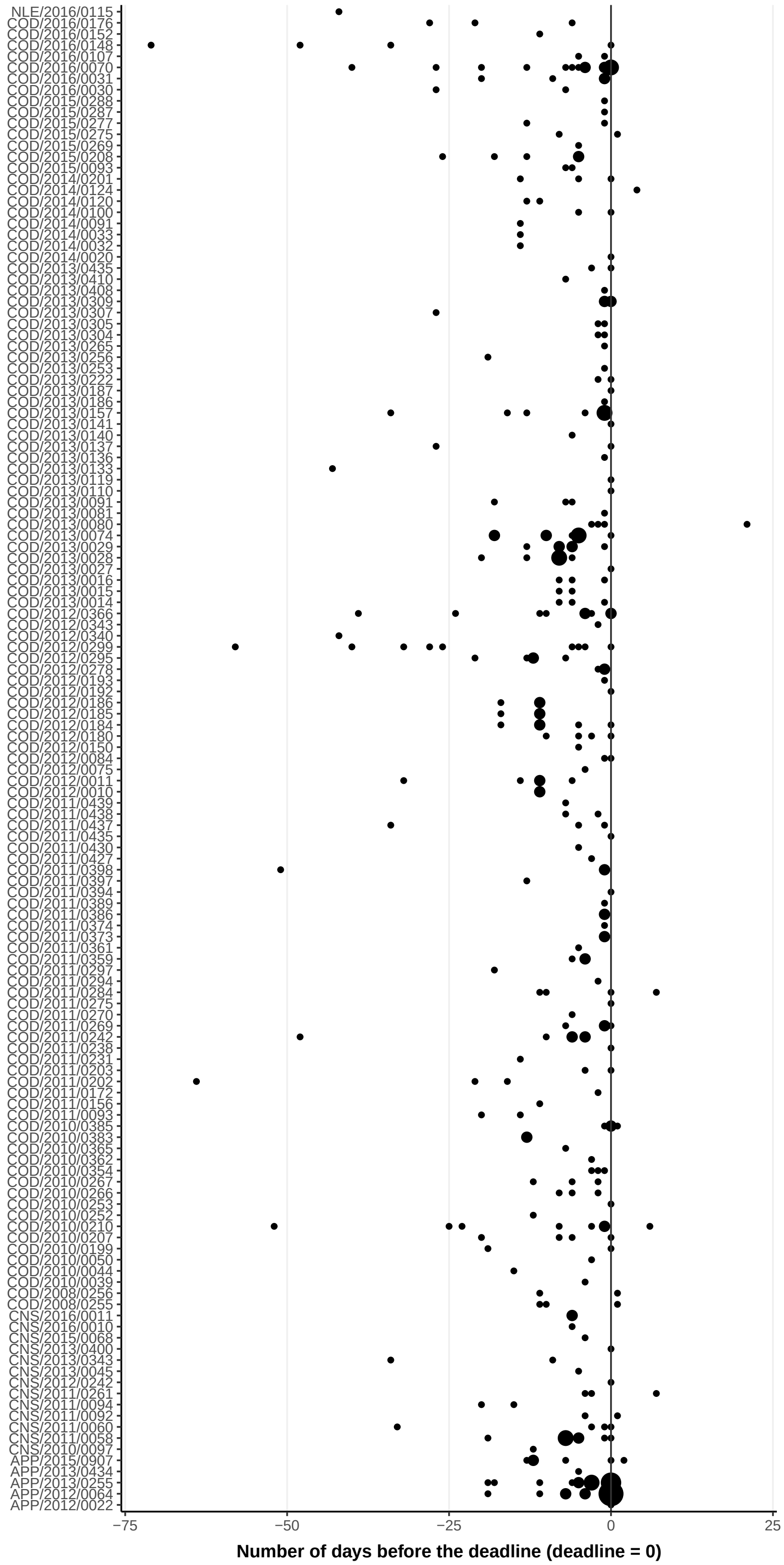


Figure 1: Timing of vetoes. Dots indicate how many days before the deadline chambers issued a veto. Dot size is proportional to the number of events occurring on the same day.

| | 1 | 2 | 3 | 4 | 5 | 6 | 7 | 8 | 9 | 10 | 11 | 12 |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| 1: Ideological homophily | | | | | | | | | | | | |
| 2: EU accession homophily | 0.46 | | | | | | | | | | | |
| 3: EU location homophily | 0.53 | 0.62 | | | | | | | | | | |
| 4: Institutional homophily | 0.49 | 0.51 | 0.49 | | | | | | | | | |
| 5: Capacity | −0.03 | −0.04 | −0.01 | 0.03 | | | | | | | | |
| 6: Control | 0.16 | 0.09 | 0.13 | 0.09 | −0.11 | | | | | | | |
| 7: EU opposition | 0.02 | 0.04 | −0.03 | 0.11 | 0.02 | 0.06 | | | | | | |
| 8: Abs. diff. in GDP | 0.36 | 0.20 | 0.22 | 0.40 | −0.02 | 0.03 | 0.01 | | | | | |
| 9: Constant GDP per capita | −0.12 | −0.14 | −0.11 | −0.10 | −0.05 | −0.36 | −0.01 | −0.04 | | | | |
| 10: Population (log) | −0.02 | −0.04 | 0.02 | −0.03 | 0.63 | 0.02 | −0.07 | −0.09 | −0.28 | | | |
| 11: Mean import from indirect ties | 0.23 | 0.38 | 0.40 | 0.51 | 0.22 | 0.00 | 0.14 | 0.23 | 0.01 | 0.17 | | |
| 12: Chamber activity | −0.07 | −0.11 | −0.10 | −0.10 | 0.06 | 0.05 | −0.14 | −0.09 | 0.28 | −0.05 | −0.08 | |
| 13: Issue specificity | −0.04 | −0.05 | −0.06 | −0.06 | 0.02 | 0.04 | −0.08 | −0.05 | 0.13 | −0.06 | −0.06 | 0.41 |

Table 1: Pairwise correlation matrix. Excessive correlation among zero values may distort the overall correlation between endogenous homophily variables.

| variable | type | level | nobs | mean/ percentage | sd | min | max | operationalization |
|---|---|---|---|---|---|---|---|---|
| Second chamber | constant | chamber | 38 | 31.60% | | | | 0/1 dummy for second chambers |
| Capacity | constant | chamber | 38 | 0.39 | 0.31 | 0.09 | 1.20 | number of seats per chamber |
| Control | constant | chamber | 38 | 1.71 | 0.58 | 0.33 | 2.67 | level of control rights (Winzen 2012) |
| EU opposition | time-varying[a] | chamber | 38 | 0.26 | 0.49 | 0.00 | 1.44 | EU negative mentions (Volkens et al. 2016) |
| Salience | constant | proposal | 140 | 8.60% | | | | 0/1 dummy for proposals on agri-culture |
| GDP per capita | time-varying (year)[b] | country | 28 | 32.00 | 21.00 | 7.00 | 102.00 | constant 2010 USDollar |
| Population | time-varying (year)[b] | country | 28 | 18145.00 | 23454.00 | 423.00 | 82133.00 | total population |
| Party family | time-varying[a] | chamber | 38 | | | | | Party family of ruling party (Volkens et al. 2016) |
| Social democratic parties | | | 9 | 23.70% | | | | |
| Liberal parties | | | 5 | 13.20% | | | | |
| Christian democratic parties | | | 6 | 15.80% | | | | |
| Conservative parties | | | 15 | 39.50% | | | | |
| Ethnic and regional parties | | | 1 | 2.60% | | | | |
| EU entry round | constant | country | 28 | | | | | Year country joined the EU |
| 1957 | | | 6 | 21.40% | | | | |
| 1973 | | | 3 | 10.70% | | | | |
| 1981 | | | 1 | 3.60% | | | | |
| 1986 | | | 2 | 7.10% | | | | |
| 1995 | | | 3 | 10.70% | | | | |
| 2004 | | | 10 | 35.70% | | | | |
| 2007 | | | 2 | 7.10% | | | | |
| 2013 | | | 1 | 3.60% | | | | |
| Political system | time-varying*c* | country | 28 | | | | | Armingeon et al. (2016) |
| Parliamentary system | | | 15 | 53.60% | | | | |
| Presidential system | | | 2 | 7.10% | | | | |
| Semi-presidential dominated by parliament | | | 11 | 39.30% | | | | |

[a] EU opposition and party family are time-varying variables. Values reported in this table represent average values of the ruling party with the longest ruling period between 2010-2016 for each chamber. [b] GDP per capita and population means are reported here for the year 2013. [c] The political system of Czechia changed from parliamentary system to semi-presidential system dominated by parliament in 2013. Percentage values reported here represent Czechia as a parliamentary system.

Table 2: Summary statistics of control variables

| | full data | without outlier | interaction effects |
|---|---|---|---|
| Ideological homophily | $0.26\ (0.09)^{**}$ | $0.39\ (0.10)^{***}$ | $0.47\ (0.11)^{***}$ |
| Ideological homophily × Proposal APP/2012/0064 | | | $-0.01\ (0.40)$ |
| Ideological homophily × Proposal APP/2013/0255 | | | $-0.20\ (0.31)$ |
| Ideological homophily × Proposal COD/2016/0070 | | | $-0.60\ (0.18)^{***}$ |
| EU accession homophily | $0.17\ (0.10)$ | $0.21\ (0.12)$ | $0.23\ (0.10)^{*}$ |
| EU location homophily | $-0.06\ (0.11)$ | $0.17\ (0.14)$ | $-0.02\ (0.12)$ |
| Institutional homophily | $0.26\ (0.06)^{***}$ | $0.31\ (0.07)^{***}$ | $0.29\ (0.06)^{***}$ |
| Abs. diff. in GDP | $0.29\ (0.04)^{***}$ | $0.26\ (0.04)^{***}$ | $0.26\ (0.04)^{***}$ |
| Second chamber | $0.59\ (0.19)^{**}$ | $0.55\ (0.20)^{**}$ | $0.68\ (0.20)^{***}$ |
| Capacity | $-0.73\ (0.49)$ | $-0.99\ (0.53)$ | $-1.05\ (0.51)^{*}$ |
| Control | $0.53\ (0.33)$ | $0.72\ (0.35)^{*}$ | $0.82\ (0.33)^{*}$ |
| EU opposition | $0.06\ (0.13)$ | $0.05\ (0.13)$ | $0.02\ (0.13)$ |
| Constant GDP per capita | $0.01\ (0.01)$ | $0.02\ (0.01)$ | $0.02\ (0.01)$ |
| Population (log) | $-0.00\ (0.13)$ | $0.06\ (0.14)$ | $0.06\ (0.13)$ |
| Mean import from indirect ties | $0.01\ (0.00)^{**}$ | $0.00\ (0.00)$ | $0.01\ (0.00)^{*}$ |
| Chamber activity | $-0.03\ (0.04)$ | $-0.03\ (0.04)$ | $-0.04\ (0.04)$ |
| Issue specificity | $0.32\ (0.09)^{***}$ | $0.32\ (0.09)^{***}$ | $0.33\ (0.09)^{***}$ |
| Salience: DG Agriculture | $-0.31\ (0.33)$ | $-0.35\ (0.34)$ | $-0.38\ (0.33)$ |
| Party family baseline: Social-Democratic | | | |
| Socialist parties | $-0.81\ (0.75)$ | $-1.16\ (0.92)$ | $-0.73\ (0.76)$ |
| Liberal parties | $-0.08\ (0.27)$ | $-0.32\ (0.29)$ | $-0.29\ (0.28)$ |
| Christian-Democratic parties | $0.03\ (0.33)$ | $-0.10\ (0.36)$ | $-0.23\ (0.34)$ |
| Conservative parties | $0.05\ (0.24)$ | $0.17\ (0.25)$ | $0.15\ (0.25)$ |
| Nationalist parties | $1.77\ (0.63)^{**}$ | $2.42\ (0.72)^{***}$ | $2.06\ (0.66)^{**}$ |
| Ethnic and regional parties | $-0.17\ (0.76)$ | $-0.13\ (0.79)$ | $0.08\ (0.77)$ |
| Entry round baseline: 1957 | | | |
| 1973 | $-0.50\ (0.32)$ | $-0.48\ (0.34)$ | $-0.40\ (0.32)$ |
| 1981 | $1.74\ (0.95)$ | $2.11\ (0.99)^{*}$ | $2.49\ (0.97)^{*}$ |
| 1986 | $0.88\ (0.62)$ | $0.93\ (0.66)$ | $1.20\ (0.63)$ |
| 1995 | $0.34\ (0.32)$ | $0.15\ (0.33)$ | $0.24\ (0.32)$ |
| 2004 | $0.10\ (0.61)$ | $0.19\ (0.66)$ | $0.50\ (0.62)$ |
| 2007 and 2013 | $-0.22\ (0.73)$ | $-0.25\ (0.77)$ | $0.17\ (0.75)$ |
| Political system baseline: Parliamentary | | | |
| Presidential | $0.24\ (0.41)$ | $0.09\ (0.44)$ | $0.30\ (0.42)$ |
| Semi-presidential dominated by parliament | $-0.02\ (0.30)$ | $-0.03\ (0.32)$ | $-0.14\ (0.30)$ |
| Proposal APP/2012/0064 | | | $-1.00\ (0.60)$ |
| Proposal APP/2013/0255 | | | $-1.08\ (0.63)$ |
| Proposal COD/2016/0070 | | | $-1.88\ (0.63)^{**}$ |
| AIC | 1420.25 | 1332.25 | 1398.76 |
| $R^2$ | 0.07 | 0.08 | 0.08 |
| Max. $R^2$ | 0.36 | 0.35 | 0.36 |
| Num. events | 353 | 339 | 353 |

$^{***}p < 0.001$, $^{**}p < 0.01$, $^{*}p < 0.05$; Endogenous network variables are rescaled (see footnote in Table 1)

Table 3: Results of the conditional logit regression on issued vetoes including interaction effects for the three yellow cards

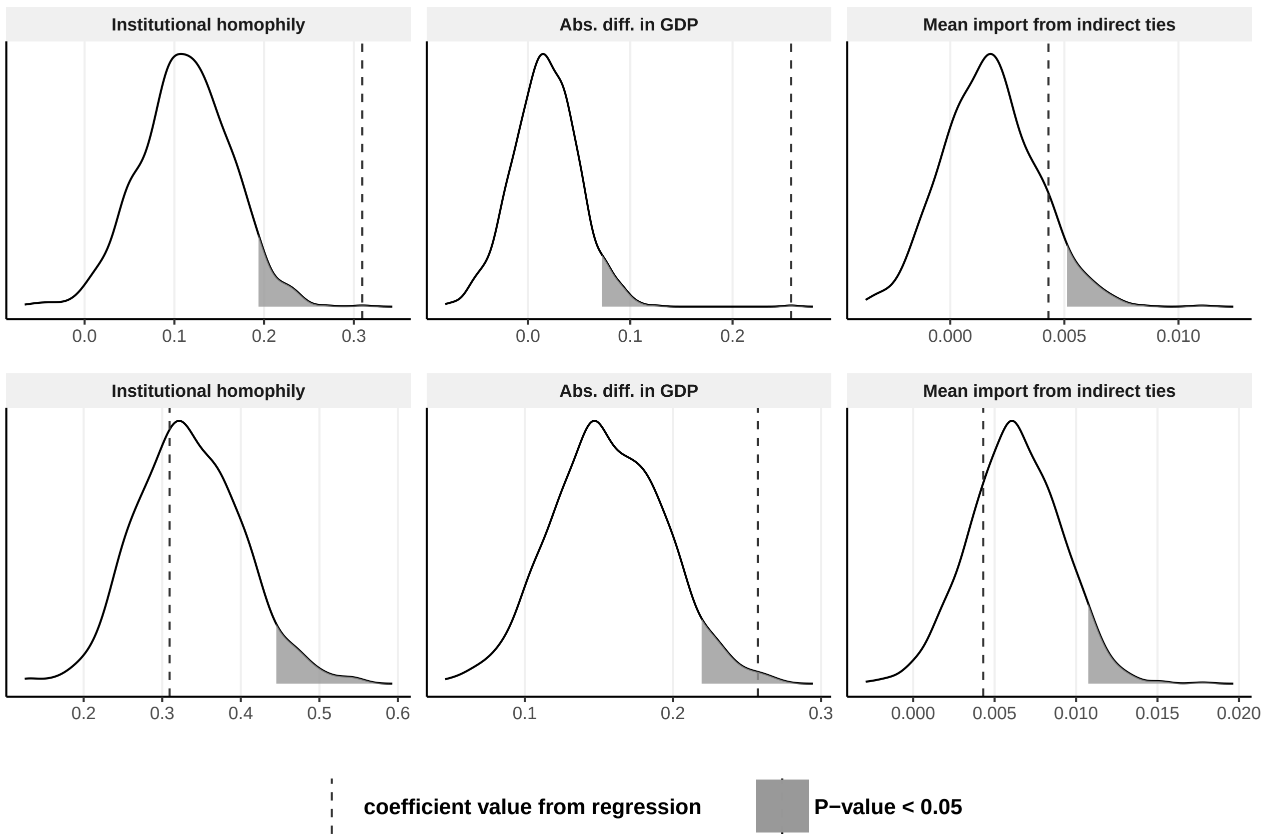


Figure 2: Permutation results for three selected controls: Distribution of coefficients across 1,000 permutations. Upper row: Dates of reasoned opinions are randomly assigned within the allotted period of deliberation. Lower row: Date assignment probability is weighted according to the time-to-event distribution found in the aggregated data (see Figure 2). The size of the unpermuted coefficient is represented by dashed vertical lines.

| | Full model | Reduced model |
|---|---|---|
| Ideological homophily | 0.39 (0.10)*** | |
| EU accession homophily | 0.21 (0.12) | |
| EU location homophily | 0.17 (0.14) | |
| Institutional homophily | 0.31 (0.07)*** | |
| Abs. diff. in GDP | 0.26 (0.04)*** | |
| Second chamber | 0.55 (0.20)** | 0.38 (0.18)* |
| Capacity | −0.99 (0.53) | −0.15 (0.48) |
| Control | 0.72 (0.35)* | 0.71 (0.31)* |
| EU opposition | 0.05 (0.13) | 0.10 (0.12) |
| Constant GDP per capita | 0.02 (0.01) | 0.02 (0.01) |
| Population (log) | 0.06 (0.14) | −0.00 (0.12) |
| Mean import from indirect ties | 0.00 (0.00) | |
| Chamber activity | −0.03 (0.04) | |
| Issue specificity | 0.32 (0.09)*** | |
| Salience: DG Agriculture | −0.35 (0.34) | −0.17 (0.28) |
| Party family baseline: Social-Democratic | | |
| Socialist | −1.16 (0.92) | −0.70 (0.81) |
| Liberal | −0.32 (0.29) | −0.15 (0.27) |
| Christian-Democratic | −0.10 (0.36) | 0.08 (0.33) |
| Conservative | 0.17 (0.25) | −0.09 (0.23) |
| Nationalist | 2.42 (0.72)*** | 2.05 (0.66)** |
| Ethnic and regional | −0.13 (0.79) | 0.32 (0.69) |
| Entry round baseline: 1957 | | |
| 1973 | −0.48 (0.34) | −0.28 (0.29) |
| 1981 | 2.11 (0.99)* | 1.04 (0.93) |
| 1986 | 0.93 (0.66) | 0.63 (0.59) |
| 1995 | 0.15 (0.33) | −0.12 (0.25) |
| 2004 | 0.19 (0.66) | 0.44 (0.60) |
| 2007 and 2013 | −0.25 (0.77) | 0.27 (0.67) |
| Political system baseline: Parliamentary | | |
| Presidential | 0.09 (0.44) | 0.01 (0.40) |
| Semi-presidential dominated by parliament | −0.03 (0.32) | −0.14 (0.29) |
| AIC | 1332.25 | 1588.31 |
| $R^2$ | 0.08 | 0.01 |
| Max. $R^2$ | 0.35 | 0.35 |
| McFadden pseudo $R^2$ | 0.19 | 0.02 |
| Num. events | 339 | 339 |

$^{***}p < 0.001$, $^{**}p < 0.01$, $^{*}p < 0.05$; Endogenous network variables are rescaled (see footnote in Table 1)

Table 4: Results of the conditional logit regression on issued vetoes comparing the full model with a model without endogenous network variables

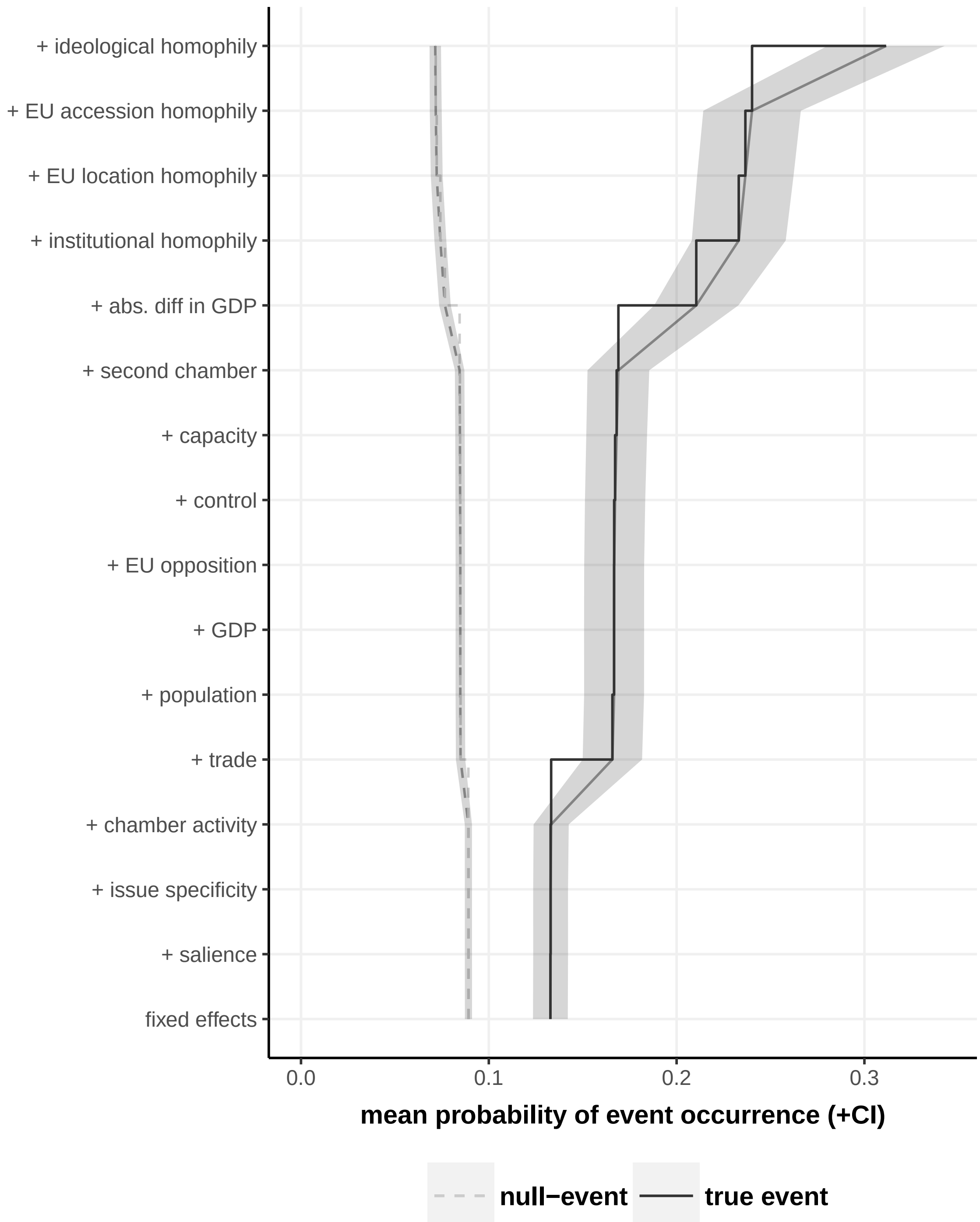


Figure 3: Stepwise model improvement measured by predicted probability. Mean predicted probability of event occurrence is shown for true events (blue) and non-events (red). 95%-confidence intervals are shown in gray. Trade, absolute difference in GDP and ideological homophily show the greatest model improvements.

| | Full model | Reduced model |
|---|---|---|
| Edges | $-8.35\ (1.39)^{***}$ | $-7.16\ (0.24)^{***}$ |
| Primary and secondary hypotheses: | | |
| Chamber clustering | $0.37\ (0.02)^{***}$ | $0.37\ (0.02)^{***}$ |
| Ideological homophily | $1.20\ (0.21)^{***}$ | $1.17\ (0.21)^{***}$ |
| EU accession homophily | $1.15\ (0.23)^{***}$ | $1.20\ (0.23)^{***}$ |
| EU location homophily | $0.39\ (0.32)$ | $0.53\ (0.26)^{*}$ |
| Institutional homophily | $1.33\ (0.20)^{***}$ | $1.32\ (0.19)^{***}$ |
| Control variables: | | |
| Second chamber | $-0.13\ (0.18)$ | |
| Capacity | $-0.20\ (0.40)$ | |
| Control | $-0.24\ (0.20)$ | |
| EU opposition | $0.06\ (0.17)$ | |
| Abs. diff. in GDP | $-8.66\ (4.70)^{\cdot}$ | $-6.85\ (4.38)$ |
| Constant GDP per capita | $15.14\ (10.48)$ | |
| Population (log) | $0.09\ (0.11)$ | |
| Mean trade with indirect ties | $0.00\ (0.00)$ | |
| Proposal clustering | $0.07\ (0.01)^{***}$ | $0.07\ (0.01)^{***}$ |
| Share of indirect ties with same DG | $3.75\ (0.41)^{***}$ | $3.75\ (0.41)^{***}$ |
| DG Agriculture | $0.47\ (0.23)^{*}$ | $0.43\ (0.23)^{\cdot}$ |
| Party family (baseline Social-Democratic): | | |
| Socialist | $0.34\ (0.45)$ | $0.28\ (0.44)$ |
| Liberal parties | $0.54\ (0.28)^{\cdot}$ | $0.46\ (0.25)^{\cdot}$ |
| Christian democratic parties | $0.34\ (0.29)$ | $0.54\ (0.23)^{*}$ |
| Conservative parties | $0.04\ (0.25)$ | $-0.14\ (0.22)$ |
| Ethnic and regional parties | $-1.57\ (0.77)^{*}$ | $-1.43\ (0.75)^{\cdot}$ |
| Entry round (baseline 1957): | | |
| 1973 | $0.45\ (0.31)$ | $0.32\ (0.24)$ |
| 1981 | $-0.04\ (0.77)$ | $-0.38\ (0.64)$ |
| 1986 | $0.56\ (0.47)$ | $0.44\ (0.33)$ |
| 1995 | $-0.05\ (0.36)$ | $-0.30\ (0.32)$ |
| 2004 | $0.43\ (0.47)$ | $-0.13\ (0.22)$ |
| 2007 and 2013 | $0.74\ (0.64)$ | $-0.15\ (0.33)$ |
| Political system (baseline Parliamentary): | | |
| Presidential | $0.43\ (0.29)$ | $0.56\ (0.24)^{*}$ |
| Semi-presidential | $0.62\ (0.26)^{*}$ | $0.44\ (0.22)^{*}$ |
| AIC | 2741.68 | 2734.93 |
| BIC | 2992.27 | 2927.05 |
| Log Likelihood | $-1340.84$ | $-1344.46$ |

$^{***}p < 0.001$, $^{**}p < 0.01$, $^{*}p < 0.05$, $^{\cdot}p < 0.1$; Coefficients can be interpreted as log-odds; Estimation of the ERGM was performed using Maximum Pseudo-Likelihood Estimation (MPLE). This may cause standard errors to be downward-biased in the model. However, the model serves well as a robustness check and the estimates are similar to the estimates in the REM.

Table 5: Bipartite ERGM of the two-mode veto network between January 2010 and September 2016

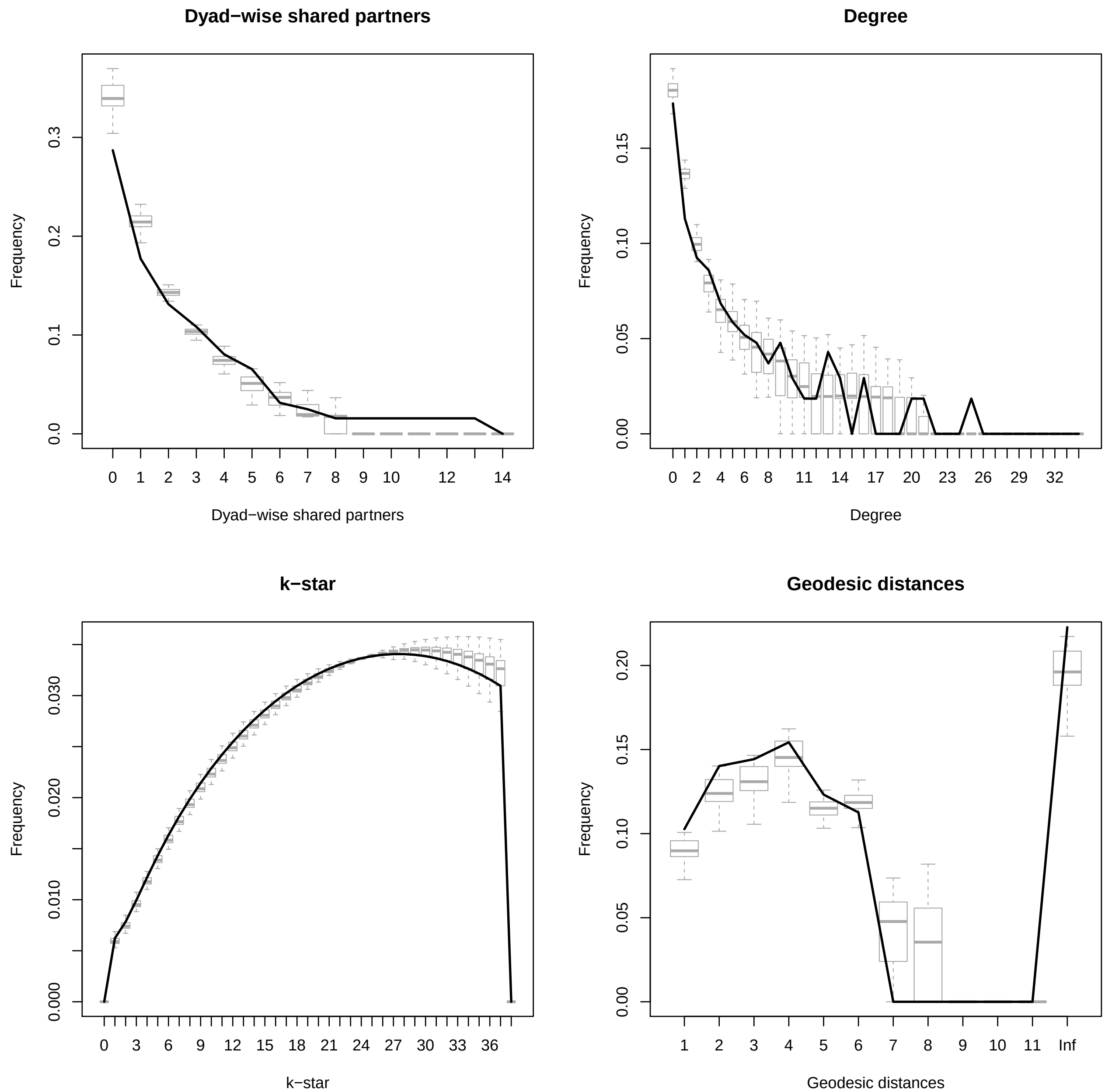


Figure 4: Goodness-of-fit assessment for the full model. The $y$-axis is log-transformed to display the nuances more clearly.

# References

Armingeon, Klaus, Christian Isler, Laura Knöpfel, David Weisstanner and Sarah Engler. 2016. "Comparative Political Data Set 1960-2014.".

Auel, Katrin, Olivier Rozenberg and Angela Tacea. 2015. "To Scrutinise or Not to Scrutinise? Explaining Variation in EU-Related Activities in National Parliaments." *West European Politics* 38(2):282–304.

Bailer, Stefanie, Mikko Mattila and Gerald Schneider. 2015. "Money Makes the EU Go Round: The Objective Foundations of Conflict in the Council of Ministers." *JCMS: Journal of Common Market Studies* 53(3):437–456.

Bennett, Colin J. 1991. "What is Policy Convergence and What Causes it?" *British Journal of Political Science* 21(02):215–233.

Buuren, Stef and Karin Groothuis-Oudshoorn. 2011. "mice: Multivariate Imputation by Chained Equations in R." *Journal of Statistical Software* 45(3):1–67.

Cranmer, Skyler J. and Bruce A. Desmarais. 2011. "Inferential Network Analysis with Exponential Random Graph Models." *Political Analysis* 19(1):66–86.

Gattermann, Katjana and Claudia Hefftler. 2015. "Beyond Institutional Capacity: Political Motivation and Parliamentary Behaviour in the Early Warning System." *West European Politics* 38(2):305–334.

Kleine, Mareike. 2013. *Informal Governance in the European Union: How Governments Make International Organizations Work*. Ithaca: Cornell University Press.

McFadden, Daniel. 1978. "Quantitative Methods for Analysing Travel Behaviour of Individuals: Some Recent Developments." pp. 279–318.

Neuhold, Christine and Alexander Strelkow. 2012. "New Opportunity Structure for the 'Unusual Suspects'? Implications of the Early Warning System for the Role of National Parliaments with the EU System of Governance." *OPAL Online Papers* 4.

Tjur, Tue. 2009. "Coefficients of Determination in Logistic Regression Models – A New Proposal: The Coefficient of Discrimination." *The American Statistician* 63(4):366–372.

Volkens, Andrea, Pola Lehmann, Theres Matthiess, Nicolas Merz, Sven Regel and Annika Werner. 2016. *The Manifesto Data Collection. Manifesto Project (MRG/CMP/MARPOR). Version 2016a.* Berlin: Wissenschaftszentrum Berlin für Sozialforschung.

Winzen, Thomas. 2012. "National Parliamentary Control of European Union Affairs: A Cross-National and Longitudinal Comparison." *West European Politics* 35(3):657–672.