Explaining Policy Ties in Presidential Congresses: A Network Analysis of Bill Initiation Data

Eduardo Alemán

University of Houston

Ernesto Calvo

University of Maryland

Policy networks formed by co-authoring and co-sponsoring bills reflect one of the most important types of connection legislators develop while in office. We expect that in presidential countries, the probability of a tie between two legislators should be influenced by partisan membership, territorial linkages and the policy areas in which they develop expertise. Given the complex nature of relational data and the particular characteristics of bill initiation networks, we propose a new approach – bootstrapping an exponential graph model using augmented data reflective of the frequency of ties – to address the challenges of *thinning* dense networks.

Keywords: legislative politics; political parties; social networks; committees

Social science literature has typically argued that dense social networks with cross-cutting affiliations provide a favorable environment for democracy and consensual politics (Lipset and Rokkan, 1967; Mutz, 2002; Putnam, 1993). Policy networks derived from co-authoring and co-sponsoring bills reflect one of the most important types of connection legislators develop while in office (Crisp, Escobar-Lemmon *et al.*, 2004). Prior works focused on bill initiation data have examined how an actor's relative position in the network affects legislative success (Fowler, 2006; Tam Cho and Fowler, 2010), how variations in ties reflect changes in political polarization (Alemán, 2009; Zhang *et al.*, 2008), how networks can be utilized to map actors' policy preferences (Alemán *et al.*, 2009; Crisp, Kanthak *et al.*, 2004) or to identify programmatic leaders (Panning, 1983). While the growing literature on social networks has illuminated some of the partisan and career traits that determine legislators' behavior, there is still relatively little comparative research on the determinants of policy networks. This article attempts to fill this gap by examining the main determinants of policy collaboration networks in two presidential democracies. Given the complex nature of relational data and the particular characteristics of bill initiation networks, the statistical analysis of policy networks raises some interesting methodological questions that our article also attempts to answer.

Policy networks are formed by purposive legislators, who connect with others with whom they share some commonality of policy preferences as well as an interest in similar policy areas or jurisdictions. These connections, we hypothesize, are well explained by the institutions structuring legislative behavior in representative democracies. A vast literature has hypothesized that electoral incentives explain both policy attention and policy effort. This motivation, characterized by David Mayhew (1974) as an 'electoral connection', is the cornerstone of legislative accountability in both candidate-centric and party-centric systems. Competing principals in the district and the nation impact legislators' collaborative efforts (Carey, 2007). The relative influence of each is supposed to stem from electoral incentives embedded in the institutional rules that regulate candidate nomination and party

competition, and the decentralization of political power. Inside congress, the organization of legislative business, particularly the committee system, further structures legislators' policy connections.

The empirical analysis presented in this article measures the determinants of policy collaboration among legislators in Argentina and Chile. The analysis of policy networks allows us to evaluate how partisan, territorial and jurisdictional determinants of collaboration operate in different institutional contexts – Argentina has a federal constitution, with a volatile party environment, and a weakly specialized committee system, while Chile has a unitary constitution with strong parties, stable coalitions and a serious committee system. Through the analysis of policy networks in these two countries we are able to illustrate within-party cohesion and cross-partisan similarities without single-mindedly focusing on floor voting behavior. The results of our analysis provide evidence of partisan, territorial and jurisdictional effects on policy design that have not been measured in prior comparative research.

The rest of this article is organized in the following manner. The first section introduces the notion of a policy network and briefly discusses the motivations behind co-sponsorship. The second presents our main argument about the formation of policy networks. The third discusses some implications for legislative behavior in the congresses of Argentina and Chile, and advances six testable hypotheses. The fourth addresses some methodological challenges facing network analyses of bill initiation data. We propose a bootstrapped exponential random graph model (B-ERGM) using augmented data reflective of the frequency of ties to deal with the problem of thinning networks. Such an approach is used to test for party, territorial and committee effects. The fifth section discusses the results from the empirical analysis, followed by our conclusion.

Legislative Behavior: What Explains the Likelihood of a Policy Tie?

A social network is composed of actors and relations between these actors. We define policy networks as the social networks formed by legislators during the process of initiating bills, and the co-authoring or co-sponsoring of a bill by two legislators as a policy tie. The relations that make up this network are joint public stances on policy which are explained by partisan and electoral linkages between representatives and their local and/or national-level constituencies. Bill initiation data, like voting or plenary floor speeches, carry information about the policy stances of legislators.[1]

What motivates legislators to co-sponsor law initiatives? The legislative literature tends to agree that co-authoring and co-sponsorship activities are carried out by legislators who are attentive, and seek to claim credit, for the policy consequences of their actions. Yet disagreement persists over the main audience for this behavior. While some scholars argue that legislators co-sponsor bills to signal their policy preferences to voters, others contend that legislators' co-sponsorship decisions are primarily a signal to fellow party members. The former perspective sees co-sponsoring as a position-taking device (Balla and Nemacheck, 2000; Campbell, 1982; Koger, 2003; Mayhew, 1974).[2] Co-sponsorship conveys information to voters regardless of floor consideration or legislative approval of initiatives.[3] In co-sponsoring, as 'in a large class of legislative undertakings, the electoral payoff is for positions not effects' (Mayhew, 1974, 132).

Several works examining patterns of co-sponsorship in the US Congress have found evidence that electorally induced position-taking influences co-sponsorship patterns (Balla and Nemacheck, 2000; Highton and Rocca, 2005; Koger, 2003; Goodliffe *et al.*, 2005; Rocca and Gordon, 2010). The literature on legislative politics in Latin America has also pointed out that legislators advertise their lawmaking activities and policy proposals to key constituencies back home. Brian Crisp, Escobar-Lemmon *et al.*, 2004, for example, presents evidence from Latin America that shows how electoral incentives facing individual legislators influence the content of the proposals they initiate in congress. A similar perspective is advanced by Crisp, Kanthak *et al.*, 2004 in their analysis of co-sponsorship in the Chilean Congress.[4]

An alternative perspective portrays co-sponsorship as a signaling mechanism with fellow legislators as the primary audience (Burkett and Skvoretz, 2001; Gilligan and Krehbiel, 1994; Kessler and Krehbiel, 1996; Wawro, 2000; Wilson and Young, 1997). The proponents of a *legislative connection* argue that legislators are motivated by policy, and seek to build ties with fellow members to facilitate the passage of bills. Consequently, legislators' co-sponsorship activities build reputations and transmit information to pivotal legislative players about the content of bills (Kessler and Krehbiel, 1996). Rick Wilson and Cheryl Young (1997), for instance, argue that US legislators co-sponsor bills as a way of signaling support, ideological content and expertise. Similarly, Daniel Kessler and Keith Krehbiel (1996) examine the relationship between legislators' preferences and the timing of co-sponsorship, finding some support for models of intra-legislative signaling.[5]

The signaling perspective predicts a different pattern of co-sponsorship ties for successful and unsuccessful bills. While Kessler and Krehbiel (1996) concentrate on testing implications related to timing of co-sponsorship and the frequency with which extremists co-sponsor, another testable implication from this theoretical perspective is that patterns of co-sponsorship derived from bills passed (i.e., a small subsample of all the co-sponsorship data) should reflect much stronger linkages among ideological opposites and much weaker linkages among legislators with similar ideological positions than in the analysis derived from the entire sample of co-sponsorship ties.[6]

While these two perspectives disagree on the targets of the co-sponsorship signal, consensus exists that the co-authoring and co-sponsoring of legislation convey policy content and reveal preferences for particular proposals *vis-à-vis* the current status quo (Grant and Pellegrini, 1998; Krehbiel, 1995; Mayhew, 1974; Talbert and Potoski, 2002). As noted by Krehbiel, 'legislators cosponsor measures whose anticipated policy consequences they like relative to the status quo, and they choose not to cosponsor measures whose anticipated policy consequences they dislike relative to the status quo' (Krehbiel, 1995, p. 910).

A small number of existing studies of co-sponsoring and co-authoring in Argentina and Chile have also characterized ties as evidence of commonality in policy preferences (Alemán, 2009; Alemán *et al.*, 2009; Crisp *et al.*, 2005). Less attention, however, has been given to the relative weight or importance of three key determinants of the co-sponsorship activity: the party, the district and the policy area or jurisdiction. In this article we assess the degree to which policy preferences are explained by partisan, territorial and jurisdictional linkages. In the next section we expand on the sources promoting bill connections between legislators.

Parties, Districts and Policy Jurisdictions

Like theorists in both the position-taking and the signaling camps, we assume that legislators' co-sponsorship connections capture their support for some policy alternative over the status quo. Like theorists working within the position-taking perspective, we assume that potential career benefits are an important determinant of co-sponsorship behavior. This may occur because legislators genuinely carry interests similar to those of their constituents or because their career goals lead them to advocate such positions. In congress, legislators should tend to connect with other legislators who hold similar views about what constitutes good public policy. Policy networks should reflect this association.

In addition, we expect policy networks to reflect legislators' interest over similar policy domains. Since bill initiation is a voluntary endeavor that communicates both preference and attention to specific issues, bill initiation behavior differs from other legislative strategies such as voting, where legislators are compelled to cast votes on (some of) those bills reaching the floor. In co-sponsoring, legislators are selective about which bills to endorse; after all, they only co-sponsor a very small sample of all the initiated bills.

Lastly, we expect policy networks to reflect legislators' opportunities for communication and coordination. Communication allows for information to flow and improves legislators' ability to evaluate others' ideological stances and policy interests. The collective action literature, through experiments and other methods, has confirmed the intuitive notion that face-to-face communication produces substantial increases in cooperation that are maintained over time (Ostrom and Walker, 1997). Repeated personal interactions provide a favorable context for trust and reciprocity to develop.

In short, we believe that commonality of preferences, shared interest over policy areas and opportunities for interaction affect the likelihood of a policy tie. As a result, we argue, policy networks should reflect the cohesion of parties, responsiveness to district-level principals, and jurisdictional expertise. The relative importance of parties, districts and policy jurisdictions as determinants of co-sponsorship behavior, in turn, depend on the institutions and rules that structure party competition.

Political Parties

Individuals entering a political career are most often drawn into parties that advocate policies and promote interests closest to their own. Self-selection is an important source of party cohesion, not only in the US Congress, but also in most other legislatures. This is particularly true in competitive environments where parties have different and relatively well-defined public stances.

Activists and committed party voters are likely to contribute over time to the ideological cohesion of legislative parties. Activists are often motivated to punish politicians consistently advancing positions incongruent with mainstream partisan stances. Antagonizing party activists and important segments of voters by repeatedly adopting non-partisan positions is electorally risky. Thus, overcoming nomination and electoral hurdles is often harder for ideological dissenters, especially if party leaders can influence this process.

In addition to similar preferences regarding what constitutes good public policy, members of the same party also share common political goals in their organized quest for

Figure 1: Measures of Individual Preferences, Co-sponsorship and Roll-Call Votes
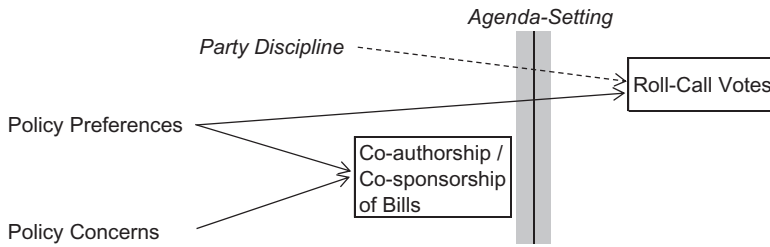
office. If we conceive of parties as groups whose members propose to collaborate in the competition for governmental power (Schlesinger, 1966; 1991; Schumpeter, 1942), then we must also consider the shared goals resulting from this collaborative effort. In the US Congress, a common interest in enhancing the value of the party label is believed to be a powerful motivator for partisan behavior (Cox and McCubbins, 2005).

Common organizational goals usually lead legislators to develop internal rules of procedure and leadership positions that help parties overcome many of the collective action dilemmas they face in congress. Leadership prerogatives, organizational procedures and legislative rules are tools that parties can use to limit the influence or shorten the careers of ideological outliers. Thus, self-selection effects are reinforced by the intervention of activists and party leaders and the constraints of institutional rules.

Common partisan membership also facilitates communication and coordination, thereby increasing the likelihood that similar preferences and interests lead to policy collaboration (Bernhard and Sulkin, 2009). There are many instances where legislators from the same party have to meet face to face, and several of these meetings are spent discussing policy issues.

For all these reasons, the likelihood of a policy tie between two randomly chosen legislators from the same party provides a very good measure of party cohesion. From a methodological standpoint, policy ties retrieved from bill initiation data are different from the information obtained through the analysis of roll-call data in three relevant aspects: (1) they tend to be unaffected by the gatekeeping behavior of congressional leaders; (2) they are less likely to be curved by whipping; and (3) they incorporate relevant information about the specific areas of policy that legislators choose to emphasize.

Because roll calls often confound partisan effects and policy orientation (both with regard to preferences and attention), as described in Figure 1, the estimates of policy commonalities differ. In countries where party discipline is very high, as is the case in Argentina and Chile, ideal points derived from roll-call votes are influenced by strong partisan effects, which tends to result in weak within-party discrimination.

Territory and Politics

Voters select their representatives from specific districts, which structure political competition. Geographical proximity often impacts on political behavior. Voters tend to be

clustered into areas with others who hold similar political preferences (Escolar, 1996; Key, 1949; O'Loughlin, 2000; Rodden, 2010). This is partly the result of economic activity, residential decisions, transportation and other types of socialization and segregation processes. It is also the result of historical–geographic processes that frame the range of political behavior in particular places (Agnew, 1996; Rodden, 2010). Territorial cleavages have been prominently discussed in existing comparative research since Seymour Martin Lipset and Stein Rokkan (1967) noted salient differences in political preferences between urban and rural constituencies of Western Europe.[7] Territorial cleavages have also been central to the political development of Eastern Europe, as shown by Joakim Ekman *et al*. (2004), and played a central role in V. O. Key's (1949) view of politics in the 'Old South', where proximity to a candidate's home town impacts on political support.[8]

Representatives elected from the same geographical areas are likely to share an interest in areas of policy relevant to local constituencies.[9] The main economic activities of the district (e.g., agriculture, fishing, mining, industry, services), its dominant social composition (e.g., urban poor, wealthy suburban, independent farms) and its geography (e.g., coastal region, desert or tropical area) make some mix of policy areas particularly relevant to citizens living in it. Legislators elected from the same electoral districts are likely to share a preference for distributive policies that target their constituents. Bills establishing targeted subsidies, funding local infrastructure projects, creating tax–free zones, providing relief to specific communities and promoting cultural events or tourist destinations are all examples of this type of activity. These bills allow legislators to claim credit for effective representation, thus members of the same district should tend to hold similar preferences with regard to such particularistic bills.[10]

Several recent works have highlighted the importance of territorial cleavages for explaining party politics in Latin America (Díaz Cayeros, 2006; Gibson and Suárez-Cao, 2010; Snyder and Samuels, 2001). With regard to Argentina, there is an abundant literature describing provincial party delegations as central to the country's democratic process. Beginning with the early studies of Gino Germani (1970) and Manuel Mora y Araujo and Ignacio Llorente (1980), an extensive literature has documented the territorial nature of political activism (Levitsky, 2003) and of party competition (Calvo and Abal Medina, 2001; Leiras, 2007) in Argentina. Such literature has shown the importance of local electoral competition in policy formation (Gibson, 1997; Gibson and Calvo, 2000), the nomination of candidates (De Luca *et al*., 2002), the allocation of patronage (Calvo and Murillo, 2004) and the distribution of social benefits (Giraudi, 2007; Weitz-Shapiro, 2008).

While there is abundant research documenting the centrality of territory for structuring party competition, there is less of a consensus about the relative importance of district-level effects for legislative behavior. With regard to the Argentine case, Mark Jones and Wonjae Hwang (2005) found no evidence that membership in the same province affects patterns of roll–call votes, while Ernesto Calvo and Andres Tow (2009) found significant province-level effects in roll–call votes pertaining to a controversial bill on the protection of native forests. In the case of contemporary Chile, which has a unitary constitution and highly nationalized parties, district-level effects have not been considered central to the interplay of politics (Magar *et al*., 1998; Navia *et al*., 2009; Siavelis, 2002; Valenzuela, 1995). Yet some works have begun to note the presence of certain territorial effects. For instance, Rodrigo Mardones

(1997) shows that legislators from the interior of the country are more likely to support decentralization policies than those coming from the city of Santiago, and Crisp, Escobar-Lemmon *et al.*, 2004 show that electoral incentives can actually foster policy collaboration with electoral rivals from the same two-member district.

Committee Work

In all presidential democracies the work of legislators is organized around a series of permanent or standing committees that have jurisdiction over well-defined areas of policy. The committee system establishes divisions of labor within congress and fosters the specialization of their members on certain issues or jurisdictions. Committee service allows legislators to develop expertise on issues which grants them authority among peers. Candidates tend to showcase this expert knowledge on the campaign trail, as they claim ownership over some policy issues and publicize their track record.

Committee assignments are, at least partly, the result of self-selection, which suggests that members of the same committee share *a priori* commonalities if not expertise. An interesting debate among US legislative scholars has addressed whether these tendencies lead to committees that are outliers *vis-à-vis* the median member of Congress. While there is relatively little evidence from roll-call votes in support of the view that US committees hold preferences significantly incongruent with the typical member of the chamber, there is some agreement that legislators have a particular interest in the policy domains of the committee in which they serve (even if this preference intensity is not consequential for voting behavior).

The recurrent contact resulting from common committee service creates many opportunities to share information about preferences and policy interests. Serving on the same committee provides a favorable context to foster the type of political friendship that makes co-authoring and co-sponsoring bills more likely. Empirical analyses of social networks have found that the work environment has an important influence on the development of friendship ties, and that ties formed among co-workers are more heterogeneous in terms of race and religion (McPherson *et al.*, 2001). With regard to legislative committees, Gregory Caldeira and Samuel Patterson (1987) argue that common service in a committee or sub-committee reflects common substantive interests and a common workload, which increases the likelihood of developing a political friendship. They confirm this hypothesis in a study of political friendship among Iowa state legislators.

There is some evidence that joint committee membership contributes to the development of policy ties. For example, J. M. Cook (2000) and Tobint Grant and Pasquale Pellegrini (1998) found evidence that in the US Congress common committee service increases the likelihood of co-sponsoring bills.

The relative importance of common committee service to the formation of policy ties can also be constrained by the institutional context. A strong committee system and candidate-centric electoral rules, for example, should increase the importance that committees have for fostering policy ties. By contrast, committees that are not highly specialized provide small returns to legislators who invest in expertise. It is unclear whether committee effects should also be present under other institutional contexts that deter specialization. In many countries, the number of committees varies significantly over time (i.e., jurisdictions

are ill defined), and the membership is considerably unstable (i.e., low incentive to specialize), which should tend to reduce the impact of shared committee membership on the likelihood of legislators developing policy ties.

To sum up, we expect committee effects, as well as partisan and territorial effects, to be present in the networks built through bill initiation. We believe these effects to be common to all presidential legislatures. In the next sections, we discuss and test these propositions with data from two South American presidential democracies.

Implications for the Study of Policy Networks in Argentina and Chile

The empirical analysis focuses on policy networks in the congresses of Argentina and Chile. Both presidential countries have had competitive democratic elections for over twenty years, and their national congresses are prominent political institutions. Since many of the insights about legislators' social networks stem from examining the case of the US Congress, it is important to see if they also apply to social networks formed in other presidential congresses. After all, generalizing in the social sciences often requires hypothesis testing using data generated from cases different from the one where the main propositions originated.

For the reasons explained in the prior discussion, we expect partisan, territorial and committee effects to have a significant impact on the likelihood that two legislators will develop a policy tie. We expect partisan effects to be strong in both Argentina and Chile. Territorial and committee effects should be present in both countries, but we expect stronger territorial effects in Argentina than Chile, and stronger committee effects in Chile than in Argentina.

The major parties of both countries are regarded as unified. In terms of roll-call votes, Argentine and Chilean parties consistently exhibit higher Rice scores than US parties. Such behavior has generally been explained by reference to party leaders' influence over legislators' electoral careers. In Chile, national party leaders, through a centralized process of nomination, veto individual candidacies and hand-pick others, while in Argentina, a decentralized nomination process has given overwhelming influence to the provincial party leadership (De Luca *et al.*, 2002). These activities not only create incentives for legislators to follow the party line when compelled to vote, but as noted before, they also contribute over time to the imposition of an upper limit on the inconsistencies of party preferences.

Argentina's largest and most dominant party since re-democratization has been the Peronist party (PJ), which usually controls between 40 per cent and 55 per cent of lower-chamber seats. The second largest party has been the historical rival of the Peronists, the Radical Civic Union (UCR). Few parties have performed as consistently as the Peronists, with a multitude of smaller parties rising and falling over short periods of time. Argentine parties tend to be poorly defined ideologically.[11] Yet legislative behavior is typically characterized as highly partisan, as in Chile, where parties are considered to have clear ideological positions.

Since the return to democracy in 1990, Chile has had two competing coalitions: the center-left *Concertación*, composed of the Socialist party (PS), the Party for Democracy (PPD), the Radical Social Democratic party (PRSD) and the Christian Democratic party (DC), and the center-right *Alianza*, composed of the Independent Democratic Union

(UDI) and National Renewal (RN). The stability of the government and opposition coalitions contrasts with Argentina, and more importantly, with the period of fluid alliances that characterized Chilean democracy prior to the military interregnum. Moreover, the reorganization of the Chilean party system in the 1980s and early 1990s led to a close association between Christian Democrats and the left, in sharp contrast to the pattern of alliances that preceded the military coup of 1973. Hence, the chances of within-coalition ties should be particularly high in Chile.

The territory of both countries suggests strong proximity effects. Both countries extend over large areas that encompass a variety of climates and regionally clustered economic activities. Argentine and Chilean regions are also strikingly different in terms of population density and wealth. Local politics are relevant in both countries. Argentina is a federal country with powerful elected governors and provincial legislatures. Chile is a unitary country, but has municipal elections and a strong historical tradition of active local politics (Valenzuela, 1977). Patterns of electoral change underscore the presence of local effects, and available evidence highlights the presence of regional partisan clustering, such as leftist support in urban and mining regions of Chile, and non-Peronist dominance in the cosmopolitan city of Buenos Aires in Argentina. This suggests that in Argentina and Chile, as in the US, voters who are geographically close to each other are more likely to hold similar political preferences. Consequently, Argentine and Chilean legislators from districts that are geographically contiguous should have constituencies with a more congruent set of political preferences than legislators from more distant districts.

The proximity aspect of territorial effects should have a strong effect in the likelihood of policy ties in both Argentina and Chile; however, two legislators from the same electoral districts should be much more likely to develop a policy tie in Argentina than in Chile. The main reason has to do with the electoral system: in general, the Chilean binomial system lessens the cooperative impact of belonging to the same district. Argentine electoral districts coincide with the jurisdictions of the 24 provinces and the federal capital (average district magnitude = 5.3), while Chilean electoral districts are small political units gerry-mandered at the end of the dictatorship to favor the exiting conservatives. While the small district magnitude in Chile (district magnitude = 2) suggest a strong incentive to attend to district interests, the *binomial* system never leads to two seats for the same party, and only infrequently leads to two seats from the same coalition. This means that the two district representatives are almost always each other's main electoral rival. This scenario is less common in Argentina where legislators are often part of provincial delegations. Moreover, many Argentine legislators pursue a political career at the provincial level after leaving congress, making district politics crucial.

Finally, we expect shared committee work to foster the likelihood of policy ties in both Argentina and Chile, but given the comparatively weaker tendencies towards specialization present in Argentina, we expect this influence to be stronger in Chile. In Argentina permanent committees are weakly professionalized, with a relatively large membership and poorly defined jurisdictions. Over the last 25 years, the number of standing committees and the number of members in each committee have both doubled. By contrast, the Chilean Congress has a smaller number of committees with well-defined jurisdictions. While the average Argentine legislator serves on five different standing committees, the average

Chilean legislator serves on two. The membership of Chilean committees is also more stable, meets more frequently and exhibits a higher degree of specialization than the membership of Argentine committees. As a result, we expect common committee membership to exert a greater effect in the Chilean Congress than in the Argentine Congress.

To sum up, the preceding discussion leads us to four main hypotheses:

> Hypothesis 1. The likelihood of a policy tie between two legislators from the same party should be significantly greater than the likelihood of a tie between two legislators from different parties.

> Hypothesis 2. The likelihood of a policy tie between two legislators from contiguous electoral districts should be significantly greater than the likelihood of a tie between two legislators from non-contiguous electoral districts.

> Hypothesis 3. The likelihood of a policy tie between two legislators from the same district should be significantly greater than the likelihood of a tie between two legislators from different districts.

> Hypothesis 4. The likelihood of a policy tie between two legislators who shared committee assignments should be significantly greater than the likelihood of a tie between two legislators who did not share such work.

Two additional hypotheses refer to specific differences we expect to find when comparing the policy networks of Argentina and Chile:

> Hypothesis 5. The effects of belonging to the same district (H3) should be strong in Argentina and weak in Chile.

> Hypothesis 6. The effects of having shared a committee assignment (H4) should be strong in Chile and weak in Argentina.

Network Analysis of Bill Initiation Data

Our analysis of policy networks uses bill initiation data from Argentina and Chile. We include all bills proposed to the Chilean House in the congressional period 2002–6 and all bill initiatives proposed in the Argentine House in the congressional period 1995–7.[12] Chilean data include information on all 493 co-authored bills initiated during the four-year term. The mean number of co-authors per bill in Chile is equal to 6.96.[13] The Argentine data are considerably larger, including 6,253 projects co-sponsored during the two-year period. The mean number of co-sponsors in Argentina is 4.95 per bill. While in Chile more legislators are linked per bill, the rate of bill initiation in Argentina is much larger.[14]

Network Analyses of Bill Initiation Data

Until recently, most statistical analyses of relational data, such as co-sponsorship data, would employ generalized linear models to estimate the parameters of interest. Such models, however, cannot properly account for the complex nature of relational data (Cranmer and Desmarais, 2009). Recent developments in the analysis of relational data have dramatically altered how we evaluate the determinants of network structure. In particular, the use of

exponential random graph models (ERGM) allows researchers to move beyond the dyadic independence assumption of earlier statistical approaches (Handcock *et al.*, 2008; Robins *et al.*, 2007).

Social networks are typically characterized by local clustering. In the prior section we advanced a series of hypotheses about the tendencies for assortative mixing (i.e., predispositions for connections with others who share certain traits) in the Argentine and Chilean policy networks. But in addition to patterns of selective mixing, studies of social networks have highlighted other propensities for tie formation that are endogenous. One mechanism that plays a prominent role in social network theories is the tendency towards triad closure: triads containing two ties will tend to form the third, creating a triangle where all three nodes are connected (Rapoport, 1957). As Steven Goodreau *et al.* (2009) note, the reasons for such a propensity towards triad closure may be due to proximity effects (e.g., two legislators connect to each other through their shared time with a third), or cognitive processes (e.g., two legislators value each other because of their agreement with a third). Thus, the process of triad closure leads to transitivity. Since assortative mixing also induces transitivity (by enhancing the chances of completed triangles within categories), an accurate assessment of the impact of one requires controlling for the impact of the other. ERGMs allow us to evaluate the impact of our propositions on the propensity for tie formation, while controlling for the propensity towards triad closure.

In an exponential random graph model a tie is assumed to be a random variable.[15] For each i and j who are distinct members of a set N of n actors, there is a random variable Y_{ij} where $Y_{ij} = 1$ if there is a network tie from actor i to actor j, and where $Y_{ij} = 0$ if there is no tie. The observed value of the variable Y_{ij} is y_{ij} while the observed network of the Y matrix of all variables is y. The probability of observing a set of network ties is formulated as:

$$P(Y = y | X) = \exp[\theta^T\, g(y, X)] / k(\theta)$$

where Y is the random set of relations in a network, y is a particular given set of relations, X is a matrix of attributes associated with the actors in the network (e.g., partisan stripe, etc.), g(y, X) is a vector of network statistics, θ is a vector of coefficients and k(θ)is a normalizing constant.[16] So the log-odds that any tie exists given the rest of the network is:

$$\text{logit}(Y_{ij} = 1) = \theta^T\, \delta[g(y, X)]_{ij}$$

where $\delta\,[g(y, X)]_{ij}$ is the change in g(y, X) when the value of y_{ij} is switched from 0 to 1.

When terms capturing endogenous effects are included, estimation is based on Markov Chain Monte Carlo Maximum Likelihood Estimation. A distribution of random graphs is simulated from a starting set of parameter values and the parameter values are subsequently refined by comparing the distribution of graphs against the observed graph with this process repeated until the parameter estimates stabilize (Hunter *et al.*, 2008; Wasserman and Robins, 2005).

Taking Advantage of the Valued Network Nature of Bill Initiation Data

In an ERGM, the socio-matrix capturing policy ties is composed of cells with dichotomous values, indicating whether each pair of legislators is connected. But as described before,

co-sponsorship information provides researchers with counts of data describing the number of dyadic ties within a given period. Different from other types of relational data such as friendship or group membership, policy networks provide information not only about the existence of a relationship but also about the relative frequency of such a relationship (i.e., the intensity of the co-sponsorship or co-authoring relation). This implies that the probability of observing policy ties differs across each pair of legislators. Because the counts observed in co-sponsorship data are meaningful, it is important to take advantage of such data rather than just assume away differences by reducing all counts to a single value expressing a tie.

Skyler Cranmer and Bruce Desmarais (2009) describe this problem as one of dealing with 'dense' networks and propose 'thinning' the co-sponsorship data by picking up a threshold to recode the network. For example, the mean number of ties among each pair of legislators in Argentina was 2.8. Co-sponsorship values above the mean could be recoded 1 while values under the threshold take the value of 0. While thinning a network using a threshold offers one way of addressing this challenge, we still want to preserve information about the different intensity or frequency of ties among pairs of legislators.

Rather than thinning the original network data using a threshold, we decided to implement a strategy that preserves meaningful information about the frequency of interaction among legislators by bootstrapping all models on simulated networks with probabilities drawn from the original data. This requires three relatively simple steps: (1) retrieve the probability matrix from the original data; (2) draw network samples from the original data; and, finally (3) run the ERGM specification on the simulated networks and retrieve mean parameters and confidence intervals.

The probability matrix for bootstrapping network graphs of bill initiation data takes as input the original affiliation matrix reporting count events and normalizes by row, dividing the count in each off-diagonal cell (representing the number of ties [i,j] between legislator i and legislator j) by the total count of bills sponsored by legislator i (the diagonal cell for each legislator [i,i]). These row-normalized data express the probability that if legislator i sponsors a bill, it will have legislator j as a co-sponsor.

We then simulate 1,000 symmetric and undirected network graphs with cell probabilities drawn from the row-normalized affiliation matrix described before.[17] Network graphs were drawn with the function *rgraph* from the library SNA in R 2.9.2, which allows researchers to provide cell-specific probabilities. An alternative to our strategy would be to simulate these networks with some parametric form derived from the original data. Because we do not want to impose any structure on the original data, we instead opted for non-parametric simulations.

Finally, we run ERGM on each of the 1,000 simulated networks.[18] The mean of the parameter estimates and confidence intervals [0.05, 0.90] are then reported. Notice that, because the probability of observing a tie among each pair of legislators differs across networks according to the frequency of their interaction, the bootstrapped estimates of ERGM carry information about the intensity of each tie. Consequently, our strategy for thinning out the data is sensitive to the actual frequencies reported in the counts of our valued network.

The Proposed Models and Variables

To test the hypotheses associated with the formation of policy networks in Argentina and Chile we estimate eight different models for each country. The dependent variables in the analysis take the value of 1 if a co-sponsorship tie is observed between each pair of legislators and 0 otherwise. As described before, we bootstrapped the original valued network to produce 1,000 simulated network graphs with dyadic ties drawn from the original probability matrix (B-ERGM). The first three models explore the effect of partisanship, territory and legislative jurisdiction on co-sponsorship networks, restricting our sample only to bills that were successfully approved on the plenary floor. The second three models explore the effect of the same covariates on the full sample of legislative initiatives proposed to the Argentine and Chilean Houses. Finally, we run two ERGM models with endogenous covariates, where we account for transitivity effects by including terms capturing the propensity towards triad closure.[19]

The model includes a series of vectors and networks as explanatory variables. First, we include variables indicating party membership. For Argentina, we include variables identifying the two main parties, PJ and UCR, and the category 'others'. For Chile we include variables for all six parties. Second, we include a variable indicating whether the legislator belongs to a government party or an opposition party. In Chile there was a multiparty coalition, as noted before, while in Argentina there was single-party government. Third, we include a variable indicating common electoral district. In addition, we include two networks as independent variables: one is a matrix capturing contiguous districts, while the second is a matrix capturing shared committee membership. We also include a control variable indicating whether the legislator is in his or her first term in office (freshman). This control is not meant to capture within-freshmen ties, but whether freshmen are more active than non-freshmen, as some of the literature from the US Congress suggests.

To control for social structure in the co-sponsorship networks of Argentina and Chile, we include three endogenous terms capturing the tendencies for triad closure. First, we add a geometrically weighted edgewise shared partner distribution (GWESP) term, which captures transitive structures in network data. This endogenous term is an alternative approach to combining k-stars and triangles (Hunter, 2007; Hunter and Handcock, 2006; Snijders *et al.*, 2006).[20] The shared partner count is taken on each edge, producing a distribution of counts that is less sensitive to network degeneracy on estimation. The GWESP statistic defines a parametric form of this count distribution that gives each additional shared partner a declining positive impact on the probability of two persons forming a tie (Goodreau *et al.*, 2009). As a robustness check, we also present results using two-stars and triangles, which provide information that is similar to GWESP. When combining stars and triangles, the former provide measures of intransitive ties in network structure while the latter provide information on expansive or transitive structures.

Results

Tables 1 and 2 present the results of the statistical models. The first model includes the government and opposition variables, while the second includes party variables. The third

Table 1: Explaining Policy Ties in the Argentine Congress, 1995–7

Variables	B-ERGM successful bills (1)	B-ERGM successful bills (2)	B-ERGM successful bills (3)	B-ERGM all bills (4)	B-ERGM all bills (5)	B-ERGM all bills (6)	ERGM all bills (7)	ERGM all bills (8)
Edges	−4.36 [−4.67, −4.09]	−4.28 [−4.59, −3.99]	−3.17 [−3.47, −2.88]	−4.62 [−5.07, −4.23]	−4.53 [−4.99, −4.13]	−3.40 [−3.80, −3.02]	−4.81 (0.21)	−4.96 (0.11)
Same committee	0.81 [0.72, 0.89]	0.85 [0.76, 0.93]	0.86 [0.76, 0.94]	0.59 [0.49, 0.68]	0.58 [0.49, 0.68]	0.60 [0.50, 0.69]	0.09 (0.08)	0.07 (0.09)
Same district	0.63 [0.48, 0.75]	0.61 [0.46, 0.75]	0.61 [0.46, 0.74]	0.98 [0.87, 1.07]	0.99 [0.88, 1.09]	0.99 [0.88, 1.09]	1.22 (0.09)	1.42 (0.06)
Contiguous district	1.33 [0.77, 1.85]	1.35 [0.78, 1.87]	1.38 [0.81, 1.89]	1.49 [0.85, 2.04]	1.48 [0.83, 2.03]	1.52 [0.86, 2.07]	0.38 (0.23)	0.24 (0.12)
Freshman	0.22 [0.16, 0.27]	0.20 [0.147, 0.258]	0.20 [0.146, 0.259]	0.06 [−0.01, 0.13]	0.07 [−0.01, 0.15]	0.06 [−0.01, 0.14]	−0.05 (0.06)	−0.06 (0.06)
Government	1.20 [1.08, 1.28]			1.22 [1.11, 1.32]			1.31 (0.11)	1.28 (0.11)
Opposition	1.19 [1.09, 1.30]			1.37 [1.26, 1.50]			1.58 (0.11)	1.51 (0.11)
Party PJ		1.11 [0.99, 1.19]			1.13 [1.03, 1.23]			
PJ with UCR			−1.26 [−1.39, −1.10]			−1.36 [−1.51, −1.19]		
PJ with other parties			−1.12 [−1.30, −0.97]			−1.08 [−1.21, −0.95]		
Party UCR		1.25 [1.12, 1.42]	0.13 [0.01, 0.30]		1.77 [1.67, 1.95]	0.66 [0.51, 0.78]		
UCR with other parties			−0.78 [−0.96, −0.63]			−0.81 [−1.00, −0.66]		
Other parties		1.95 [1.814, 2.16]	0.84 [0.76, 0.94]		1.79 [1.63, 1.93]	0.48 [0.53, 0.77]		
GWESP							−0.02 (0.02)	
Triangle								1.63 (0.00)
K-star (2)								−0.16 (0.00)
AIC	7,393	7,287	7,278	9,586	9,601	9,586	6,523	6,413
BIC	7,449	7,351	7,358	9,671	9,669	9,671	6,590	6,489
n	43,681	43,681	43,681	43,681	43,681	43,681	43,681	43,681

Table 2: Explaining Policy Ties in the Chilean Congress, 2002–6

Variables	B-ERGM successful bills (1)	B-ERGM successful bills (2)	B-ERGM successful bills (3)	B-ERGM all bills (4)	B-ERGM all bills (5)	B-ERGM all bills (6)	ERGM all bills (7)	ERGM all bills (8)
Edges	-3.57 [-3.92, -3.21]	-3.35 [-3.69, -3.02]	-3.32 [-3.71, -2.93]	-3.58 [-4.07, -3.16]	-3.15 [-3.58, -2.77]	-3.39 [-3.97, -2.92]	-4.16 (0.33)	-3.83 (0.14)
Same committee	0.57 [0.44, 0.66]	0.59 [0.46, 0.70]	0.57 [0.45, 0.67]	0.35 [0.23, 0.47]	0.40 [0.25, 0.53]	0.36 [0.23, 0.48]	0.27 (0.10)	0.34 (0.11)
Same district	0.59 [-0.01, 1.10]	0.48 [-0.07, 0.99]	0.63 [0.03, 1.14]	-0.13 [-0.93, 0.98]	-0.34 [-1.07, 0.73]	-0.09 [-0.87, 1.03]	0.59 (0.14)	0.00 (0.15)
Contiguous district	0.43 [0.27, 0.54]	0.43 [0.26, 0.54]	0.42 [0.25, 0.53]	0.23 [0.07, 0.39]	0.22 [0.05, 0.38]	0.23 [0.08, 0.39]	0.99 (0.48)	1.07 (0.49)
Freshman	0.11 [0.03, 0.19]	0.11 [0.03, 0.18]	0.11 [0.04, 0.20]	0.09 [0.01, 0.19]	0.06 [-0.02, 0.17]	0.09 [0.01, 0.19]	0.02 (0.07)	0.02 (0.08)
Government	0.81 [0.68, 0.95]			1.39 [1.24, 1.58]			1.37 (0.14)	1.16 (0.13)
Opposition	1.35 [1.22, 1.48]			1.58 [1.38, 1.77]			1.52 (0.14)	0.68 (0.17)
Party PS		1.59 [1.15, 2.08]			1.66 [1.07, 2.10]			
Party PPD		-0.08 [-0.38, 0.29]			1.39 [1.07, 1.65]			
Party PRSD		0.37 [0.09, 0.77]			1.03 [0.06, 1.94]			
Party DC		1.13 [0.89, 1.32]	1.10 [0.83, 1.35]		1.28 [1.05, 1.53]	1.51 [1.14, 1.84]		
Party RN		1.85 [1.65, 2.01]			1.61 [1.37, 1.86]			
Party UDI		1.65 [1.43, 1.83]			1.85 [1.65, 2.04]			
Leftist bloc (PS-PPD-PRSD)			0.54 [0.33, 0.79]			1.37 [1.07, 1.65]		
Leftist bloc with DC			0.36 [0.17, 0.58]			0.90 [0.69, 1.22]		
Leftist with rightist			-0.45 [-0.67, -0.21]			-0.36 [-0.62, -0.02]		
Rightist bloc (RN-UDI)			1.10 [0.93, 1.27]			1.38 [1.15, 1.59]		
GWESP							0.03 (0.03)	
Triangle								-0.09 (0.00)
K-star (2)								0.13 (0.00)
AIC	3,682	3,572	3,668	3,953	3,868	3,938	3,002	3,001
BIC	3,731	3,648	3,737	4,001	3,944	4,007	3,064	3,639
n	14,641	14,641	14,641	14,641	14,641	14,641	14,641	14,641

is designed to capture tendencies for cross-partisan ties. The first three are run with the smaller sample of successful bills, and the last two are ERGM models that include endogenous terms.

Overall, the results lend strong support to our main hypotheses. As expected, legislators are very likely to connect with other legislators with whom they share partisan labels, territory and committee assignments. These effects are significant regardless of whether we restrict our sample only to successful bills, or include endogenous controls.

In both Argentina and Chile, government (opposition) legislators are significantly more likely to develop ties with other government (opposition) legislators than with members of the opposition (government). Similarly, legislators are significantly more likely to develop within-partisan ties than cross-partisan ties. For example, the probability of a tie between two government legislators in Argentina is around 3.3 times higher than the likelihood of a tie between a government legislator and an opposition legislator.[21] Within-party ties are also significantly more likely to develop than cross-partisan ties, for both the PJ and the UCR as well as for the 'others' category (models 2 and 5). In Chile, the probability of a tie between two government (*Concertación*) legislators is also much higher than the likelihood of a tie between a government legislator and an opposition legislator (4 times higher when considering all bills, and 2.2 times higher when only considering successful bills). Partisan ties are also significantly more likely than cross-partisan ties (except for the PPD when considering only successful bills). In both countries, these effects remain significant whether we consider bills approved on the plenary or all law initiatives, and after controlling for the endogenous tendency towards triad closure (model 7) and two-stars and triangles (model 8).

The results highlight differences in partisan cohesion. For example, while all Argentine legislators show a tendency toward intra-partisan connections, legislators from the UCR appear almost twice as likely to develop within-party ties as Peronist legislators (model 6). This reflects the divergence of positions within Peronism that is usually highlighted in the qualitative literature. The most cohesive Chilean party appears to be the rightist UDI, and the least cohesive ones seem to be the small leftist parties, PRSD and PPD: UDI legislators are 5 times more likely to develop within-party ties than cross-partisan ties, while PRSD and PPD legislators are, respectively, 2.8 and 4 times more likely to develop within-party ties than cross-partisan ties (model 5). But overall, the cohesiveness of the rightist block (UDI-RN) is no different from the cohesiveness of the leftist parties (PRSD-PPD-PS) within the *Concertación* coalition (model 6).

The results highlight patterns of cross-partisan commonalities. In Argentina, small parties appear more likely to develop policy ties with the UCR than with Peronist legislators (model 3). Still, the least likely policy tie is between a UCR and a PJ legislator. In Chile, the model capturing cross-group ties differentiates between leftists (social democrats in PPD and PRSD, and socialists in PS), Christian Democrats (DC) and rightists (RN and UDI). Since the left out category in this model is a tie between DC and rightists, the results show that when all bills are considered, the odds of a DC legislator developing ties with a leftist legislator is 2.5 times greater than the odds of developing ties with a rightist legislator (model 6).[22] During this congressional period, the least likely connection in the Chilean Congress is between leftist and rightist legislators.

The results also convey the relevance of territorial effects. In Argentina, a policy tie between two legislators coming from the same province and from provinces that are contiguous is significantly more likely to develop than a policy tie between two legislators who do not share these territorial traits. The impact of contiguous province appears as the strongest predictor in the Argentine network. The odds of connecting are 2.7 times greater for legislators from the same province and 4.4 times higher for legislators from contiguous districts than for those not sharing these territorial traits (model 4). These effects are robust to all the model specifications. While territorial effects are also present in Chile, they are less consequential than in Argentina. The effect of contiguous district is positive and significant in all models, while the effect of belonging to the same district is only significant in models 3 and 7. For example, the odds of a tie between two legislators from contiguous districts are 1.3 times higher than those between two legislators not sharing any territorial traits (model 4).

Committee effects are evident in both countries: two legislators sharing committee assignments are significantly more likely to develop a policy tie than two legislators who have not shared such work. For Argentina, the coefficient is highly significant and surprisingly strong in all six bootstrapped models. It is revealing that, in both countries, committee effects are consistently stronger in the sample of successful bills than in the full sample. While this finding deserves further scrutiny in future work, one tentative explanation for the strong effect of committees in Argentina is that common *a priori* interest in similar areas of policy (manifested by similar assignments) and the interaction of committee work, even if limited, are sufficiently potent to foster collaboration. Thus, the division of labor intrinsic to legislatures, the committee system, promotes the development of policy ties, even if individual members lack the expertise developed through high tenure and professionalization.

Finally, the results for the control variables reveal other interesting results. The coefficient for first-time legislators captures the tendency towards tie formation, not a tendency for within-group ties, as with the other variables. The results show that in the Chilean policy network, first-time legislators show a slight tendency to co-author more bills than more senior legislators. In the Argentine policy network, the effect is significant only when analyzing the restricted sample of passed bills. In both Chile and Argentina the GWESP estimates failed to achieve statistical significance.[23] Triangles and two-stars in Argentina and Chile, on the other hand, are shown to be statistically significant and describe different tendencies in the co-sponsorship networks. That is, while in Argentina legislators from the same party show a tendency to form horizontal ties with fellow representatives (tendency to co-sponsor with friends-of-a-friend); in Chile negative and statistically significant two-stars reflect a more hierarchical within-party network.

To sum up, the results are consistent with most of our hypotheses. The likelihood of two legislators developing policy ties is significantly higher if these legislators are from the same party, from contiguous electoral districts or have been assigned to work on the same congressional committee. As expected, belonging to the same electoral district has a major impact on the likelihood of developing a policy tie in Argentina, and not much effect under the two-member districts used in Chile. Unexpected, however, were the results for shared committee work in Argentina, where a weakly institutionalized committee system had

suggested a weaker effect. This implies that even in the context of weakly institutionalized committee systems, shared policy interests and opportunities for interacting can make policy ties between committee members highly likely. The results also informed us about the relative cohesion of parties, the unity of governing coalitions and the likelihood of cross-partisan ties. In Chile, the tendency towards cross-partisan ties follows coalition attachments and a clear ideological ordering. In contrast, the pattern of cross-partisan ties in Argentina highlights the division between the two major parties and contenders for the presidency.

Conclusion

This article provides a number of theoretical, methodological and empirical insights. We begin making the case that institutional incentives (partisan, territorial and organizational) influence the formation of policy ties. We argue that value networks carry relevant information about tie formation, and offer a strategy to weight ERGM estimates by the frequency of observed ties, with few assumptions and without imposing a parametric form that may bias our estimates. The article also contributes to the empirical analysis of legislative behavior by providing the first cross-national analysis of policy networks (and their determinants) in Latin America. The results offer clear evidence of systematic influences on legislative behavior and measure the different partisan, territorial and organizational determinants of policy ties. In contrast to prior research in Argentina, which has had difficulties assessing the importance of provincial politics in the policy-making process, we show territoriality to be one of the most important determinants of policy affinity. This finding complements other non-legislative research that has deemed inter-province coalitions critical for explaining voting patterns in Argentina (Calvo and Escolar, 2005; Calvo and Murillo, 2004; Gibson, 1997). Results from the Chilean policy network confirm the findings from prior works showing high levels of coalition unity and an ideological alignment of parties (Alemán and Saiegh, 2007; Carey, 2002), and provide new insights into the legislative consequences of committee work and territorial proximity.

To conclude, the study of social networks derived from bill initiation data offers one important, yet seldom utilized approach to learn about legislators' policy positions. For scholars interested in comparative legislative analysis, the greater availability of bill initiation data with respect to roll-call data should be welcome news.

About the Authors

Eduardo Alemán is an Associate Professor in the Department of Political Science at the University of Houston. He specializes in the comparative analysis of political institutions and Latin American politics. His current work focuses on executive–legislative relations, parliamentary procedures, government coalitions and agenda control in Latin American legislatures. Professor Alemán has published articles in such journals as *World Politics*, *Comparative Politics*, *Comparative Political Studies*, *Legislative Studies Quarterly* and *Latin American Research Review*. Eduardo Alemán, Department of Political Science, University of Houston, 447 Philip G. Hoffman Hall, Houston, TX 77204-3011, USA; email: *ealeman2@uh.edu*

Ernesto Calvo is an Associate Professor of Political Science at the University of Maryland. His research focuses on issues of political representation, elections and congresses. His research, which has received awards from the Comparative Politics sections of the American Political Science Association and the Latin American Studies Association, has

been published in US, European and Latin American journals such as the *American Journal of Political Science*, the *Journal of Politics*, *World Politics*, the *British Journal of Political Science*, *Comparative Political Studies*, *Electoral Studies*, *Política y Gobierno* (Mexico), *Desarrollo Económico* (Argentina), *Opiniaõ Publica* (Brazil) and the *Revista de Ciencia Politica* (Chile). Ernesto Calvo, University of Maryland, Government and Politics, 3144F Tydings Hall, College Park, MD 20742, USA; email: *ecalvo@gvpt.umd.edu*

Notes

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 1 Co-sponsorship data have distinctive advantages such as the use of a large sample of bills and no sample selection bias, given that information is available for all initiatives proposed to congress. Co-sponsorship data also tend to be more informative than roll-call data with regard to positive ties. However, the decision not to co-sponsor a bill may reflect opposition to a bill but it may also reflect lack of interest or lack of information. Because ties (the ones in the data) are highly informative but non-ties (the zeros in the data) are not, most analyses of co-sponsorship data treat differently the decision to co-sponsor bills from the decision not to co-sponsor a bill (Alemán *et al.*, 2009).

 2 Mayhew (1974) saw co-sponsoring as involving few costs and potentially large benefits, but others disagree about the low cost. For instance, Koger (2003) notes how co-sponsoring can cost political support as well as earn it. Rocca and Gordon (2010) argue that the chance of voter retaliation exists because there are a variety of intermediaries with sufficient resources to inform constituents. Schiller (1995), as well as Rocca and Gordon (2010), highlight that co-sponsorship often involves costly time, and Fowler (2006) notes that the members of the US Congress only co-sponsor about 3 per cent of proposed bills and that bill authors spent much effort seeking co-sponsors.

 3 In the US Congress close to 7 per cent of bills initiated by legislators become law, in the Argentine Congress the figure is close to 6 per cent, while in the Chilean Congress it is just over 16 per cent.

 4 Kessler and Krehbiel (1996, p. 555) also predict 'a close correspondence between legislators' ideological predispositions (or, by extension, those of their reelection constituencies) and the content of the legislation they choose to cosponsor'. Several works link the ideological positions of US legislators to their co-sponsorship decisions (Grant and Pellegrini, 1998; Highton and Rocca, 2005; Panning, 1983; Regens, 1989). Moreover, Alemán *et al.* (2009) show that ideal points derived from roll-call votes are highly correlated with ideal points derived from co-sponsorship data.

 5 See Goodliffe *et al.* (2005) for a recent challenge to their findings.

 6 A test of these two competing models is beyond the scope of this article. Data, code and ancillary information comparing both sets of models are available from: http://www.bsos.umd.edu/gvpt/calvo/

 7 See also Rokkan *et al.*, 1983.

 8 Gimpel *et al.* (2008) argue that this advantage is rooted in the link between physical proximity and familiarity, which induces trust. See also Chen and Rodden, 2009.

 9 For evidence linking co-sponsorship activities to territory, see Caldeira and Patterson, 1987; Cranmer and Desmarais, 2011; Grant and Pellegrini, 1998; Regens, 1989.

10 Territorial effects should be contingent on the electoral rules in place. See Crisp, Escobar-Lemmon *et al.*, 2004.

11 Very high turnover rates in the Argentine Congress have led to a membership characterized by what Jones *et al.* (2002) call professional politicians and amateur legislators.

12 While reporting the results for all congressional periods falls beyond the scope of this article, we find that the estimates of partisan, district and committee effects are roughly similar in magnitude and significance to those reported for the period 1995–7.

13 The mean number of co-sponsors is similar to the one observed in the US House, although the Chilean House has only a quarter as many members (Alemán *et al.*, 2009).

14 In the affiliation matrix we observe higher counts in Argentina (2.87) than in Chile (1.94). The density of the Argentine policy network is 0.369 (i.e., the proportion of all possible ties present), while the density of the Chilean policy network is 0.547.

15 This summary of ERGM modeling is based on Handcock *et al.*, 2008; Robins *et al.*, 2007.

16 See Goodreau *et al.*, 2009, 7–8.

17 To ensure that the network is symmetric and takes advantage of the information contained in the affiliation matrix, we simulate 500 network graphs from the lower triangle and 500 network graphs from the upper triangle of the original affiliation matrix.

18 We use the statnet package for all ERGM analyses (Handcock *et al.*, 2008).

19 These last two models are run with the binary networks indicating the presence or absence of a tie (i.e., they do not incorporate information about the frequency of ties).

20 Alternative models and replication material are available from: http://www.bsos.umd.edu/gvpt/calvo/

21 The interpretation of the coefficient is in terms of log-odds. So, for example, the log-odds of a tie between two Chilean legislators who do not share any traits (and are not freshmen) is given by the edges coefficient, which is −3.58 in model 4 (i.e., the probability is 3 per cent).

22 This legislator would be 1.4 times more likely to develop such ties if we restrict the analysis to successful bills only.

23 The values we used for GWESP were 0.7 for Argentina and 0.25 for Chile. These models were stable. Others showed GWESP as significant but also with greater instability. Our main hypotheses, however, remained robust to these changes in GWESP.

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