What Does it Take for Congress to Enact Good Policies? Policy-specific Information and Roll-Call Voting in the U.S. Congress

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Abstract

We study the conditions under which members of Congress incorporate policy-specific considerations in their voting decisions. To do this, we estimate a model that accounts for uncertainty and private information about legislation quality, identifying the sources of heterogeneity in responsiveness to policy-relevant information. We show that legislators' electoral concerns and institutional position strongly influence the likelihood of evaluating initiatives on their merits. In particular, uncompetitive House races and low incumbent turnover are detrimental for information-based voting. Through their impact on representatives' decisions, these factors aggregate and transmit information about bill quality across both chambers of Congress, ultimately affecting policy outcomes.

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1 Introduction

Modern pieces of legislation are complex objects, putting forth elaborate solutions to multiple intertwined issues. This is true for both "technical" legislation away from the public eye and heavily publicized bills alike, be it health care reform, financial regulation, or immigration. In this context, the oftentimes useful analytical simplification of left and right-wing politics falls short of capturing some of the key aspects of the decision-making problem faced by members of Congress (MCs).

Consider, for example, the *Americans with Disabilities Act* (ADA) of 1990. While the main goal of the bill was to improve the working conditions for disabled employees, Acemoglu and Angrist (2001) found that due to the additional costs imposed on employers the ADA actually reduced job opportunities for young disabled workers. This is a bad outcome for all legislators, left or right. Similar stories of good intentions undone by perverse incentives reemerge in the *Endangered Species Act* of 1973 (List, Margolis and Osgood, 2006) or the *Wild and Free-Roaming Horses Act* of 1971 (Winerip, 2013). In other instances, still, unintended or failed policies arise for reasons other than badly designed economic incentives, due to the incorrect assessment of the environment in which the law takes effect (e.g., *Iraq War Resolution*). In fact, in many issues, the apparent "ideological" divisions are at heart disagreements about the relative effectiveness of alternative policies to attain some common objective, based on limited information available to politicians or to society in general.

The point is that an integral part of the production of legislation is the assessment of objective relations between policies and the environment in which these policies take effect, many of which are hard to pin down precisely. Getting these objective relations right is what we refer to as quality. What does it take for Congress to enact high quality legislation? Under what conditions will representatives incorporate policy-specific information at the moment of casting a vote in Congress?

Surprisingly, we know relatively little about this. While political scientists have long recognized that bringing about good public policy is one of the main goals of members of Congress (Fenno, 1973; Kingdon, 1977), most of the empirical congressional literature focused on purely ideological or distributional problems, disregarding the quality dimension of legislation. Even the work that addressed quality head on - most notably Krehbiel (1991) and Epstein and O'Halloran (1999) - concentrated primarily on its ramifications for the institutional organization of Congress. The contributions in this area were mostly theoretical, centering on the repercussions of electoral considerations and career concerns on the incentives of elected politicians to vote informatively or pander to the public (e.g., Maskin and Tirole, 2004; Canes-Wrone and Shotts, 2007). However, we know from the voluminous research on the U.S. Congress that, besides individual preferences, institutions *and* electoral concerns matter for voting decisions. We argue here that these factors, together with characteristics of the proposals under consideration, also matter to determine the conditions under which representatives pursue high quality legislation.

A crucial empirical hurdle is that it is difficult to measure the quality of policy in a systematic and comparable way. Iaryczower, Katz and Saiegh (2013) bridge this gap by formulating a model in which members of Congress are imperfectly informed about the quality of legislative proposals. Their model builds fundamentally on the bicameral nature of the U.S. Congress and formally develops the intuitive notion that, if legislators have private information about the relative quality of the alternatives, voting outcomes in the originating chamber can aggregate and transmit information to members of the receiving chamber. By structurally estimating the model, the authors can recover bill quality and the proportion of legislators who vote based on the merits of the proposals in a similar manner as to how Poole and Rosenthal (1985) recover legislators' ideal points.

In this paper we extend Iaryczower, Katz and Saiegh (2013)'s approach, quantifying the role of bill-specific characteristics, electoral considerations and aspects of the internal organization of Congress on legislators' propensity to incorporate policy-relevant information in their decisions. In order to do this, we exploit equilibrium information to recover the parameters linking predictive factors of interest to voting behavior, resorting to developments in latent class regression analysis (Ungar and Foster, 1998; Huang and Bandeen-Roche, 2004) to fit the model *via* Markov chain Monte Carlo simulations.

Our results show that a large fraction of House members (40.8%) base their voting decisions on the quality of the alternatives under considerations. Characteristics of the initiative such as whether it was publicly supported by the President and the proportion of minority party members cosponsoring it provide MCs with prior information about their quality. The extent to which representatives act on such information is in turn contingent on their personal characteristics. In particular, legislators' electoral concerns and position in the internal organization of Congress have a first order effect on their proclivity to rely on private information at the moment of casting a vote on the floor.

Leadership, seniority and committee membership, all have a significant impact on the number of representatives voting informatively, and thus on information aggregation and the pursuit of high quality legislation. In line with our expectations, we find that majority party leaders are more likely to vote unconditionally in favor of proposals than the rank and file, while MCs with no gate-keeping power and no stake in getting particular pieces of legislation through Congress are more prone to decide based on their own judgment about the quality of bills. The main result, however, concerns the impact of seniority. We find that minority party newcomers are roughly twice as likely to vote informatively as more experienced members of the opposition. An increase in the number of terms served by majority party members, on the other hand, has a positive but marginal impact on their propensity to decide based on policy-relevant information. The net effect of an increase in seniority is therefore to reduce the prevalence of informative voting in the House: whereas in a chamber composed exclusively of long-term incumbents only a third of the MCs would vote based on the perceived quality of legislation, the proportion would rise to 43% if all representatives were freshmen.

Electoral safety and, to a lesser extent, constituents' knowledge of their representatives and their policy positions are important predictors of informative voting as well. Compared to minority legislators elected in close races, members of the opposition facing little or no competition in their districts are less inclined to vote based on their private information and more predisposed to oppose proposals independently of their quality. Even when the margin of victory does not significantly affect information-responsiveness among representatives in the majority, each percentage point increase in the vote-share difference between the winner and runner-up across all congressional districts would reduce the average probability of informative voting by 0.12 points.

Taken together, the results for seniority and electoral competition have strong implications for institutional design and the quality of democratic representation. Our estimates indicate that reducing actual and potential renewal of the membership leads to a 20 percentage point decline in the use of information in policy-making. This finding is particularly troublesome in view of the historically high incumbent reelection rates in the U.S. Congress and the trend towards less competitive districts verified in the last decades (Campbell and Jurek, 2003). Furthermore, since we show that non-competitive races and low House turnover also have a negative impact on the probability that bills originating in the lower chamber are approved in the Senate, these factors ultimately undermine the quality of legislation enacted by Congress. Our analysis thus provides strong empirical support for theoretical arguments suggesting that term limits and reforms aimed at increasing electoral competitiveness may contribute to improve public policy (e.g. Maskin and Tirole, 2004).

2 Related Literature

The goal of this paper is to determine which factors influence MCs' propensity to incorporate policy-specific information in their voting decisions. To the best of our knowledge, our article provides the first empirical analysis of this issue. In doing so, however, we draw on the contributions of a large literature.

Two influential works, Krehbiel (1991) and Londregan (2000), are closely related to our research. Krehbiel (1991) initiated a prolific literature on the informational role of congressional committees. The starting point for this work is the existence and relevance of what we call a quality dimension in legislative policy-making. But while in our model information is dispersed across all members of Congress, Krehbiel studies the transmission of information between the median member on the committee – who is informed about the realization of a policy-relevant state – and the median member on the floor, who is not. Additionally, while our analysis examines the impact of legislator, bill-specific and contextual factors on voting strategies and policy outcomes, Krehbiel considers the repercussions of the theory for the organization of Congress.

Londregan (2000) introduces *valence* in an empirical model of legislative policy-making. As in our paper, the attention of the analysis is focalized on legislators' voting behavior. The fundamental difference with our approach is that Londregan's valence is a publicly known quality of legislation. There is no uncertainty about whether the proposal "gets the environment right", and therefore no private information. In this setting, then, it is impossible to assess which are the most important factors leading MCs to rely on their judgments about the merits of the proposal at the moment of casting a vote on the floor. In our paper, instead, the quality of each bill is unknown, and representatives are imperfectly informed about it.

More generally, our paper draws on a vast body of empirical work examining the determinants of legislators' preferences and voting behavior. Most of this research, though, is grounded in the spatial voting theory (Poole and Rosenthal, 1985, 1997). In this private values model, members of Congress are perfectly informed about the characteristics of the alternatives under consideration and simply choose the one that is closer to their ideal policy. Concerns about whether and to what extent the proposals might be a poor response for the current state of affairs do not affect MCs' decisions. Consequently, these studies have little to say about the conditions under which representatives will be more or less likely to respond to policy-relevant information.

Even analyses that do not adhere to this purely ideological account of legislative decision-making typically treat the consideration of bills in the House and the Senate as theoretically and statistically independent. The underlying assumption in these studies is that members of one chamber cannot obtain any relevant information about the initiatives by observing the voting outcomes in the other chamber. Iaryczower, Katz and Saiegh (2013), however, estimate a model of congressional decision-making that allows for dispersed information about bill quality in an equilibrium context and show that legislators in the receiving chamber can use the information conveyed by votes in the originating chamber to improve their own decisions.

The core of our empirical strategy is akin to that of Iaryczower, Katz and Saiegh (2013). As in their work, representatives in our model receive private signals from a distribution that is conditional on the realization of an unobservable or latent state variable (the quality of legislation). Unlike Iaryczower, Katz and Saiegh (2013), though, we gauge the impact of individual, contextual and bill-specific factors on MCs' disposition to evaluate proposals based on their merits. To do this, we adopt a more flexible and general econometric approach that allows us to identify and estimate the sources of heterogeneity in legislators' responsiveness to policy-specific information.

3 Theory and Empirical Model

Our voting model introduces a simplified description of a bicameral legislature to capture the core incentive problems for the transmission and aggregation of information with sequential committees. Members of Congress choose between a proposal A_t and a status quo SQ_t. The proposal is considered sequentially by the two chambers, the House (H) and the Senate (S). Chamber $j = H, S$, is composed of n_j members. The alternatives are first voted on in the House; members of the Senate observe the outcome of the vote in the originating chamber, and then vote between the two alternatives. The proposal passes in chamber j if it receives at least $(n_j + r_j)/2$ votes, for $r_j \in \{1, \ldots, n_j\}$, and is adopted by Congress if and only if it passes in both the House and the Senate.

To this basic setting we add the various components of the model: the information of the different agents (prior beliefs and private information), and their preferences. We then describe equilibrium behavior, connecting underlying parameters to voting outcomes.

Public Information. MCs are imperfectly informed about the quality of the proposal being voted in roll call t, $\omega_t \in \{0, 1\}$. Representatives cannot observe the quality of the bill, which can be high $\big(\omega_t = 1\big)$ or low $\big(\omega_t = 0\big)$, and for each roll call $t = 1, \ldots, T$ have a prior belief $p_t \in (0, 1)$ that the bill is of high quality. These beliefs are common knowledge for legislators but uncertain for the econometrician. We assume that MCs' prior beliefs that the proposal is of high quality are given by:

$$Pr(\omega_t = 1 | \mathbf{x}_t, \nu_{c[t]}, \vartheta_{g[t]}) = p_t = \frac{\exp(\mathbf{x}_t^{'}\alpha + \nu_{c[t]} + \vartheta_{g[t]})}{1 + \exp(\mathbf{x}_t^{'}\alpha + \nu_{c[t]} + \vartheta_{g[t]})} \tag{1}$$

where ν_c, ϑ_g are error terms accounting for unobserved heterogeneity in the quality of bills across Congresses and issue areas, respectively, and $\mathbf{x}_t$ includes bill-specific characteristics which we expect to be correlated with perceptions of bill quality.

One such characteristics is cosponsoring activity. Because MCs want to avoid being blamed for failed policies, they are unlikely to cosponsor bills they believe to be of low quality. As a result, the *number of cosponsors* of a bill can be a signal of quality to both the econometrician and other legislators (Woon, 2008). Raw cosponsoring data, however, can be a noisy signal of quality, since a higher number of cosponsors may be driven by ideological proximity rather than by the attributes of the proposal (Campbell, 1982). To account for this fact, we distinguish between partisan and bipartisan cosponsorship by incorporating the *proportion of minority cosponsors* among the predictors in $\mathbf{x}_t$.

Another factor likely to influence p_t is the President's decision to publicly support a legislative initiative. Both because they do not want to be associated with failed policies and because they do not want to back bills that will not pass, Presidents will tend not to support low quality legislation (Marshall and Prins, 2007). Since this is common knowledge among MCs, the *presidential position* on a proposal - coded on an ordered scale ranging from 1 for votes publicly opposed by the President to 3 for votes he backed - can be informative to members of Congress.[1] Because high quality bills are more likely to be approved by Congress in our model, we expect presidential support to be positively correlated with the quality of legislative initiatives.

A third element that can affect MCs' perception of quality is the salience of the issue. To the extent that constituents tend to be more informed about and hold representatives accountable for prominent roll call votes (Ansolabehere and Jones, 2010), salience can induce legislators to put more effort and attention to assure good quality legislation. On the other hand, salience can also be a manifestation of partisan antagonism and make it more difficult to reach agreements about policy contents (Shull and Vanderleeuw, 1987). In this case, legislators might be willing to sacrifice quality considerations in favor of

[1]Using dummy coding for *presidential position* - with "no position" as the reference category - does not change the substantive findings reported in Section 5.

ideological concerns. Which of these two possibilities prevails is therefore an empirical question. In our analysis we use Congressional Quarterly (CQ)'s definition of *key vote* as a proxy for salience. CQ classifies votes as "key" if they pertain to matters of major controversy, involve decisions of potentially great impact on the nation and lives of Americans, or are a test of presidential or political power (CQ 2006 Almanac Plus, p. C-3). Although these criteria are admittedly rough, key votes are generally recognized as the most important ones of any session (Jesse and Theriault, 2014), and relying on CQ's widely used operationalization avoids the potential arbitrariness of alternative ad-hoc definitions while maximizing the consistency of this variable throughout the period under study.[2]

Private Information. In addition to the public information contained in p_t, each MC i in the House receives an imperfectly informative signal $s_{i,t} \in \{-1, 1\}$ about the quality of each bill. Individuals' signals are i.i.d. conditional on ω_t, with $\Pr(s_{i,t} = 1|\omega_t = 1) = \Pr(s_{i,t} = -1|\omega_t = 0) = q_t > 1/2$. The precision of the signal is common knowledge for legislators but uncertain for the econometrician, and we assume that in each roll call vote t, q_t follows a normal distribution truncated in the $(1/2, 1)$ interval:

$$q_t(\mathbf{w}_t, \beta, \epsilon_{c[t]}, \varphi_{g[t]}) \sim TN(\mathbf{w}_t'\beta + \epsilon_{c[t]} + \varphi_{g[t]}, \sigma_q^2, 0.5, 1) \tag{2}$$

where φ_g and ϵ_c are random terms allowing the precision of signals to vary across policy areas and Congresses, and $\mathbf{w}_t$ includes the number of words of the bill, the number of

[2]The basic results do not change if votes identified as important by interest groups like Americans for Democratic Action or the American Conservative Union are used. However, the roll call votes included in these groups' classifications are more ideologically biased and their numbers considerably smaller than those in CQ's sample (Cox and McCubbins, 1993).

committees it was referred to, and representatives' prior experience with similar legislation as proxies for the information content and complexity of the proposal (Epstein and O'Halloran, 1999; Maskin and Tirole, 2004).

Preferences. MCs care about the quality of the bill, but also have ideological biases: a preference over proposals that is unrelated to their merits. We assume that in each roll call t, each MC i has a publicly known bias either for or against the proposal, and we say that i is *pro-change* or *anti-change* respectively. Pro-change MCs face a cost of $\pi_t^P \in (0, 1)$ if Congress approves a low quality bill and a cost of $1 - \pi_t^P$ if it does not approve a high quality proposal. Anti-change MCs, on the other hand, face a cost of $\pi_t^A \in (\pi_t^P, 1)$ if Congress approves a low quality proposal and a cost $1 - \pi_t^A$ if it does not approve a high quality bill. The payoffs for pro (anti) change MCs if Congress approves a high quality proposal (rejects a low quality bill) are normalized to zero. Thus, while all representatives prefer a high quality proposal, they differ in the amount of information supporting an alternative that would induce them to vote for it. Given information $\mathcal{I}_i$, pro-change MCs prefer the proposal to the status quo whenever $\Pr(\omega_t = 1|\mathcal{I}_i) \geq \pi_t^P$, while anti-change MCs are open to support the proposal only if $\Pr(\omega_t = 1|\mathcal{I}_i) \geq \pi_t^A$, where $\pi_t^A > \pi_t^P$.

Equilibrium Voting Behavior. We consider Perfect Bayesian equilibria in pure strategies in which at least some members of the House vote *informatively*; i.e., in favor of the bill when their private assessment is that the proposal is of high quality, and against it otherwise. We concentrate on equilibria in which only members of the House vote informatively, since in our data House bills are almost never killed on a vote in the Senate (see Section 4).[3]

[3]Equilibria in which members of both the originating and receiving chambers vote informatively require by construction that bills approved in the House pass/fail a vote in the

In all equilibria with these characteristics, members of the Senate disregard their private information and act only to raise the hurdle that the alternative has to surpass in the House to defeat the status quo, killing the bill when the vote tally in the House is below an endogenous majority rule (EMR) and approving it otherwise. In equilibrium, the endogenous majority rule in the Senate and the voting strategies of members of the House are such that legislators in both chambers have incentives to follow their equilibrium behavior. In particular, House members voting informatively have incentives to do so because, conditional on affecting the outcome in the Senate, their inference on the information of other members of the House voting informatively exactly compensates their bias.[4]

EMR voting equilibria separate members of the House in three behavioral types $\theta_i \in \Theta \equiv \{I, Y, N\}$. Each individual can be voting informatively ($\theta_i = I$), uninformatively in favor of the proposal ($\theta_i = Y$), or uninformatively against the proposal ($\theta_i = N$). Conditional on w_t, a legislator i voting informatively will support a high quality proposal with probability q_t, and a low quality proposal with probability $1 - q_t$. A legislator voting uninformatively for (against) the proposal, on the other hand, supports the proposal with probability 1 (0), independently of the state. In other words, while a representative voting informatively will decide based on her private information about the bill under consideration, a representative voting uninformatively for (against) a proposal will do so *even if* she receives negative (positive) information about its merits.

Senate with positive probability. In the kind of equilibria considered here, in contrast, it is irrelevant whether a proposal fails in the Senate because it is voted down or because it is never taken up for consideration. See Iaryczower (2008) for details.

[4]We relegate a formal statement of the results and their proof to the Supplementary Materials Appendix (Section S.1), and present here an informal description of equilibrium voting strategies.

In equilibrium, the behavioral type of each legislator is known for other MCs. However, this is uncertain for the econometrician. We assume that the probability that legislator i is a behavioral type $l \in \Theta$ is given by:

$$\Pr\big(\theta_i = l \big| \mathbf{z}_i, \eta_{c[i]}, \varepsilon_{g[i]}\big) = \frac{\exp(\mathbf{z}_i' \gamma_l + \eta_{c[i],l} + \varepsilon_{g[i],j} + \varsigma_{s[i],l}.)}{\sum_l \exp(\mathbf{z}_i' \gamma_l + \eta_{c[i],l} + \varepsilon_{g[i],l} + \varsigma_{s[i],l})} \tag{3}$$

where $\mathbf{z}_i$ is a vector of legislator- and constituency-specific variables, and $\eta_{c,l}$, $\varsigma_{s,l}$ are Congress and constituency random effects accounting for the hierarchical nature of our data and for the inclusion of district-level explanatory variables among the predictors of θ_i (Gelman and Hill, 2007).[5]

We allow MCs' tendency to respond to policy-relevant information or to back/oppose proposals on purely ideological grounds to depend on several covariates that figure prominently in the literature on the U.S. Congress. Partisanship is obviously a crucial predictor of legislators' voting behavior, potentially reflecting average differences in preferences across groups of like-minded individuals and mediating the impact of institutional and electoral factors. In a purely partisan decision-making environment, we would expect *majority party* members to be generally biased in favor of the proposals advanced by the party controlling the House (and the Rules Committee), while representatives in the minority would tend to be predisposed against these initiatives. However, if - as we argue - quality matters in Congress, party labels should not completely determine observed voting patterns.

[5]Failing to do so would violate the assumption of independent and identically distributed errors, potentially leading to negatively biased standard errors and erroneous inferences. The number of congressional districts (435) is large enough to render this multi-level specification suitable for addressing the potential intra-cluster correlation induced by the inclusion of constituency-level predictors in equation 3 (Snijders and Bosker, 2012).

Prior research has also suggested that the institutional position that legislators hold can affect whether and to what extent they pursue policy seeking - as opposed to position taking - goals (Woon, 2008). We capture this institutional dimension by MCs' standing in the legislative and partisan hierarchies, as summarized by the attainment of leadership positions and seniority status, and by their role in the consideration of the bill - more concretely, whether or not they belong to the committee in which the bill was originated.

Due to their unique role in the legislative process, we expect the majority party *leadership* to be more inclined than members of the rank and file to give their unqualified support to bills that are put up for a vote. This is for several reasons. First, the leadership will tend to advance bills that they themselves do not oppose. Second, once the leadership chooses which bills to advance, its own success or failure is determined - at least in part - by whether these bills pass or not (Sinclair, 1983). Third, majority leaders oftentimes provide a signaling function towards the rank and file with the goal of protecting the brand image presented to the electorate in congressional elections (Sinclair, 1983; Cox and McCubbins, 1993). Hence, majority leaders should be less likely to condition their voting decisions on the quality of legislation than rank and file members. For the minority leadership the expectations are less clear-cut: without agenda setting power, the gate-keeping effect is moot, as is the payoff to delivering the passage of proposals taken up for a vote.

The *seniority* of members of Congress provides a more nuanced measure of their political clout. Traditionally, the seniority system guaranteed that long-serving representatives would advance in the House hierarchy. Once they had risen to positions of power, experienced MCs could essentially disregard the wishes of the party (Congressional Quarterly, 2012). Hence, the incentives to vote uninformatively along party lines - rather than considering the merits of the proposals - should be highest among the most junior members of the House. Despite the erosion of seniority norms governing career advancement in the last decades, freshman MCs still face more uncertainty about their reelection and are thus more dependent on partisan benefits than more seasoned members of Congress, who enjoy a more established reputation and name recognition (Stratmann, 2000). Based on these arguments, we expect representatives' propensity to vote in accordance with their own private information to be positively correlated with the number of terms they serve.[6] This "liberating" effect of seniority should be stronger for the majority party, which controls most positions of power as well as the resources with which to help junior members.

Committee membership is presumably correlated with useful policy expertise and a more intimate knowledge of the characteristics of the proposals (Krehbiel, 1991; Cox and McCubbins, 1993). Hence, representatives involved in drafting the bills will tend to receive stronger private signals about legislation quality and should be in a better position to evaluate proposals on their merits than non-committee members. On the other hand, there has been much academic debate regarding whether and to what extent committees actually provide Congress with useful technical information or, on the contrary, comprise preference outliers with a stake in delivering specific legislation and getting it passed (Fenno, 1973; Kollman, 1997). In the latter case, members' biases could take precedence over their private information at the moment of casting a vote. Ultimately, which effect prevails can only be determined by the data.

The final pillar shaping legislators' responsiveness to policy-specific information is given by what we can broadly call the reelection motive (Mayhew, 1974). The theory of elections as a disciplining device suggests that, under certain conditions, electoral competition reduces public officials' willingness to stick to specific policy positions disregarding relevant information (see Maskin and Tirole, 2004, and the references

[6]To account for possible non-linearities in the relationship between the number of terms served and θ_i, we include a quadratic term for *seniority* in $\mathbf{z}_i$.

therein). In particular, reelection-seeking politicians facing a serious electoral challenge have incentives to follow their signals and choose the "correct" alternative in order to convince constituents of their ability to gather accurate policy information before the next election (Canes-Wrone, Herron and Shotts, 2001). Additionally, a central contention of the political agency literature is that, when citizens are politically uninformed and/or there is uncertainty about the policy congruence between legislators and voters, reelection concerns can induce distortions in politicians' behavior, leading them to pander to the electorate rather than to decide based on information that is not available to their constituents (Maskin and Tirole, 2004; Canes-Wrone and Shotts, 2007). Thus, other things equal, we expect MCs from safe constituencies or from districts comprised mainly of uninformed voters to be less likely to guide their decisions by their private information than representatives from more competitive or politically aware districts. We measure electoral competitiveness by the *margin of victory* between the incumbent and her closest challenger in the previous House race.[7] The degree of political knowledge of the electorate, in turn, is approximated with two correlates (Snyder and Strömberg, 2010): *lack of familiarity with the incumbent*, given by the average proportion of individuals in a district who cannot recognize their representative or are unable to place her on a seven point ideological scale; and *constituents' level of information*, defined as the average frequency with which constituents read daily newspapers.[8]

[7]As a robustness check, we also operationalized this variable with a binary indicator taking the value 1 if the incumbent won by at least 60% of the vote and 0 otherwise (Canes-Wrone, Brady and Cogan, 2002). This has little effect on our substantive findings.

[8]We also estimated our model substituting these measures with district-level socio-demographic characteristics that have been shown to be correlated with political information (e.g., education, income). The results are similar to those reported below.

4 Empirical Approach

Data. Our data consists of roll call votes, legislator, constituency and bill-specific covariates. The voting data comprise all bills that were originated in the House and whose passage was decided by a roll call vote between 1991 - 2006 (Congresses 102 through 109). A total of 818 such bills were considered for approval in a vote on passage in the House over this period.[9] Of the 778 proposals that made it out of the House, less than 1.3% (10) failed to pass a vote in the Senate. However, more than 45% (360) were never taken up for consideration on final passage in the receiving chamber, and thus failed de facto (see Figure S.1 in the Supplementary Materials Appendix).

The regressors included in the analysis follow the discussion in Section 3, in addition to several control variables commonly used in the congressional literature (e.g., indicators for divided government, election year and legislators' last period in office, the size of the House majority, the policy area the bill belongs to, the previous presidential vote in the district, and constituency-level socio-demographic characteristics). A detailed description of the coding and sources for all these variables, along with descriptive statistics, can be found in the Supplementary Materials Appendix (Section S.2).

Model identification. Formal statistical conditions for identification of finite mixtures of generalized linear models such as ours are given in Huang and Bandeen-Roche (2004) and Grün and Leisch (2008), among others.[10] Here we present a more informal argument

[9] Ideally, it would also have been desirable to include bills originated in the Senate in our analysis. However, during the period under study, there were only 106 bills originating from the Senate whose passage was decided by a roll call vote. Due to this data limitation, in this paper we restrict our attention to House bills.

[10] We verified that the necessary and sufficient conditions for identification of mixtures of binomial distributions discussed by Huang and Bandeen-Roche (2004) are met in our

regarding the identification of θ, q and p from the observed data (see also Figure S.2 in the Supplementary Materials Appendix for a graphical illustration).

Intuitively, the common value component of the model plays a key role for identifying these parameters. Suppose first, for ease of exposition, that representatives' prior beliefs about the quality of the proposals p and the precision of their private information q were invariant across bills. Assume further that p was close to 1/2, indicating a roughly fifty-fifty chance of "high"/"low" quality proposals, and that q was close to 1, meaning that legislators' private information about bill quality was very accurate. Under these conditions, representatives voting informatively would switch between virtually unanimous "yea" and "nay" votes across roll calls. If, keeping p fixed, the precision of private signals decreased, we would observe more dispersed voting patterns among these informative MCs, with some of them voting correctly (i.e., in accordance with ω) for or against the proposals and the rest voting incorrectly in the opposite direction.

Next, assume that, for a certain value of q, p increased towards 1. Then, holding the signal precision constant, MCs basing their decisions on their private information would tend to vote "yea" more often as the prior probability of "high quality" bills became larger. The frequency of informative votes for the proposals will thus track p, with higher values of the prior corresponding to a larger frequency of "yea" votes among type-I MCs.

It follows from the arguments above that if the common prior beliefs about the quality of the proposals and the precision of the private signals were both high (i.e., close to 1), nearly every informative voter would support the initiatives under consideration. As either p or q declined, MCs for which $\theta = I$ would alternate between "yea" and "nay" votes. These changes, however, would have no effect on the behavior of House members voting uninformatively: those assigned to behavioral type $Y(N)$ would - almost - always

application. Additionally, "fake data simulations" (Gelman and Hill, 2007) showed that our estimation strategy is able to recover the true model parameters.

back (oppose) the initiatives regardless of the values of p and q. Hence, low variability in individual decisions and lack of co-movement with other MCs' votes will distinguish non-informative from informative voters.

It is important to stress that, as discussed above, our empirical analysis controls for a rich set of contextual, bill- and legislator-specific covariates. Thus, MCs classified as voting informatively are those less likely to exhibit immovable voting records *after* accounting for disparities in the content and characteristics of the proposals as well as for differences among representatives due to their partisan affiliation, position in the congressional and partisan hierarchies, and electoral environment.

Estimation strategy. In order to estimate our model, we integrate developments in collaborative filtering and latent class regression analysis (Ungar and Foster, 1998; Huang and Bandeen-Roche, 2004). This approach allows us to recover the key unobservable variables (θ, ω, q) of our decision-making model from observed voting patterns, while simultaneously quantifying the influence of the predictors of interest on legislators' prior beliefs about the quality of the proposal, their information and behavior.

Estimating the impact of these predictors is relatively straightforward once MCs have been assigned to behavioral types and bills categorized as "high" or "low" quality. The problem is more involved, though, because both θ_i and ω_t are unknown, latent quantities, which are to be estimated. However, this can be achieved by using the information from our model. Note that given observed roll-call votes, $y = \big(y_1, \dots, y_T\big)$, we can write the marginal distribution $P(y)$ (ignoring random effects for ease of exposition) as:

$$P(y) = \prod_{t=1}^{T} \sum_{s\in\{0,1\}} \sum_{i\in n_H} \sum_{l\in\Theta} \Pr\big(y_{i,t}|\theta_i = l, \omega_t = s, q_t, \mathbf{X}_t\big) \Pr\big(\theta_i = l\big|\mathbf{z}_i\big) \Pr(\omega_t = 1|\mathbf{x}_t), \qquad (4)$$

where $\mathbf{X}_t \equiv (\mathbf{x}_t, \mathbf{w}_t, \mathbf{z}_i)$, $\Pr(\omega_t = 1|\mathbf{x}_t)$ is given by (1), $\Pr\big(\theta_i = l\big|\mathbf{z}_i\big)$ is given by (3), and the probability of observing a vote in favor of the proposal if $\theta_i = I$ is

$$\Pr(y_{i,t} = 1|\theta_i, \omega_t, q_t) = \begin{cases} (1-q_t)(1-\mu) + q_t\mu & \text{if } \omega_t = 0 \\ q_t(1-\mu) + (1-q_t)\mu & \text{if } \omega_t = 1 \end{cases} \tag{5}$$

while

$$\Pr(y_{i,t} = 1|\theta_i, \omega_t, q_t) = \begin{cases} 1-\mu & \text{if } \theta_i = Y \\ \mu & \text{if } \theta_i = N, \end{cases} \tag{6}$$

where μ is a probability of error such that whenever equilibrium behavior dictates a vote $v_{i,t} \in \{0,1\}$ we observe $y_{i,t} = v_{i,t}$ with probability $1-\mu$ and $y_{i,t} = 1 - v_{i,t}$ with probability μ. The introduction of μ as an additional parameter to be estimated allows relaxing the theoretical model's assumption of a deterministic relationship between legislators' types and roll call votes, accommodating other unobserved influences on individual voting patterns besides θ_i and its predictors.[11]

Our interest lies primarily in the parameters α, β, and γ of expressions (1)-(3). Because θ_i and ω_t can be seen as missing data, parameter estimates can be obtained by Markov chain Monte Carlo (MCMC) simulations that utilize data augmentation techniques for the latent variables of the model. In a nutshell, the MCMC algorithm alternates between: i) generating random draws for each θ_i and ω_t from the posterior probabilities of class membership given the observed data and parameter estimates; ii) drawing new values for the remaining parameters of the model from the augmented data posteriors which regard class-membership indicators as known. Repeating these steps generates a sequence of iterates converging to the stationary posterior distribution.

[11] Hence, MCs can in practice be classified as voting uninformatively even if they do not always support/oppose every proposal with probability 1, which is a very restrictive assumption. We estimated μ to be 0.11 across all Congresses and policy areas, but the results do not change if it is fixed at reasonably small values (e.g., between 0.05 and 0.15).

In our application we ran three parallel Markov chains with dispersed initial values for 200,000 cycles each, discarding the first half as burn-in. We fitted the model separately for key and "non-key" votes and then pooled the weighted posterior samples drawn by the MCMC algorithm to summarize the distribution of the model parameters. This allows the assignment of legislators into behavioral types to vary independently between prominent and less salient initiatives - i.e., to decide informatively in some (e.g., key) votes but uninformatively in others - while preserving model identifiability (Huang and Bandeen-Roche, 2004; Jesse and Theriault, 2014).[12] Section S.3 in the Supplementary Materials Appendix shows that our estimation approach correctly predicts about 80% of the individual decisions in our sample and outperforms alternative methods commonly used in the congressional literature.

5 Results

In this section we present our main results. Section 5.1 examines the impact of legislator, bill-specific and contextual variables on voting behavior in the House, focusing on the influence of these factors on the probability that MCs respond to policy-specific information. Section 5.2 then explores how the information about legislation quality contained in House vote tallies affects passage rates in the Senate.

[12] We also estimated a more flexible econometric specification allowing all the characteristics of the bills to influence the posterior probability of legislators' classification into types. While this specification is not strictly derived from our theoretical model, the main results are in line with those presented in Section 5.

5.1 Determinants of informative voting in the House

We begin by assessing the (unconditional) impact of partisanship on MCs' equilibrium behavior.[13] Figure 1, which contrasts the posterior distribution of behavioral types within the majority party and the opposition, shows that there is a clear partisan division in legislators' inclination to incorporate policy-specific information in their decisions. On average, almost 90% of the majority party members support any given bill unconditionally, while only 9% vote according to their assessment of the merits of the proposal. In contrast, a whopping 77% of the opposition votes in favor of the bill if and only if they have a positive private assessment of its quality, while 18% votes against the bill regardless of their beliefs about its value.

The fact that we classify a large proportion of the minority as voting informatively has a clear counterpart in the raw data. As noted by Krehbiel (1998, p. 6), winning coalitions in the House are normally much greater than minimum-majority size and typically bipartisan, both at the level of roll-call votes generally and votes on final passage more specifically. The stylized fact identified by Krehbiel is the reduced form representation of the underlying equilibrium voting strategies we estimate.

Figure 1 here

Despite the marked partisan differences in voting patterns, Figure 1 also indicates that more than 40% of the members of the House do not exhibit systematic biases for or against the proposals. In the context of our model, these are the informative voters who incorporate quality considerations in their decisions. However, it is important to take into account another possible explanation for this finding, namely, that representatives who do

[13] Table S.4 in the Supplementary Materials Appendix displays the "raw" parameter estimates. However, because these are quite difficult to interpret, we center the discussion on "auxiliary" quantities such as average predictive comparisons (Gelman and Hill, 2007).

not toe the party line might be voting in accordance with their own ideological preferences, rather than based on policy-relevant information.[14] In other words, the crucial distinction between informative and uninformative voters in our model could conceivably be driven by the existence of preference outliers conditional on party affiliation. In particular, given that the growing party polarization observed in the contemporary Congress has given rise to heightened levels of party voting (Cox and McCubbins, 1993), an obvious question is whether what we interpret as information-based decision-making is simply a synonym for preference moderation.

Figure 2 examines this issue, summarizing the relationship between MCs' behavioral types and their first-dimension DW-NOMINATE scores, typically interpreted as reflecting legislators' relative positioning on the liberal/moderate/conservative policy space (McCarty et al., 1997).[15] The left panel of the figure plots the relationship between the absolute value of representatives' scores and their posterior probability of being classified as informative voters. While the posterior mean of $P(\theta = I)$ is highest among relatively centrist MCs, the credible intervals overlap across the whole ideological spectrum, indicating no significant differences in information-responsiveness at conventional levels. Moreover, the right panel shows that the distribution of ideal points is statistically indistinguishable across behavioral types for the 102-108 Congresses. Only in the 109th

[14]We thank an anonymous referee for bringing this point to our attention.

[15]Despite some caveats, DW-NOMINATE scores remain the most common measures of congressional ideology. Since including these scores as predictors in our model would lead to obvious endogeneity problems, we only explore the bivariate correlations between them and behavioral types. For robustness, we also fitted our model including Bonica (2013)'s CFscores - based on campaign contributions - as regressors in equation 3. These estimates, reported in Figure S.6 of the Supplementary Materials Appendix, do not indicate a systematic relationship between ideology and informative voting either.

Congress do we observe a significant difference between spatially conservative type-Y legislators - most of whom belong to the Republican majority of the House - and the rest. Even in this session, though, the alignment of informative voters along the liberal-conservative continuum largely coincides with that of representatives opposing proposals irrespective of their merit. Similar conclusion are drawn if attention is restricted to majority or minority party MCs only (Figure S.7 in the Supplementary Materials Appendix).

Figure 2 here

The evidence in Figure 2 underlines that the spatial placement of type-I voters is quite spread out throughout the period under analysis, and that moderate MCs are not systematically more responsive to private information than other legislators. In fact, among representatives with first-dimension DW-NOMINATE scores in the $(-0.25, 0.25)$ interval - a range of values typically seen as comprising moderate members of Congress (McCarty et al., 1997) - the average probability of casting an informative vote is only slightly above one half (0.54). Furthermore, as shown in Figure S.8 of the Supplementary Materials Appendix, almost 40% of these moderate MCs are classified as voting uninformatively for or against proposals, and this proportion exceeded 60% in some congressional sessions (e.g., 102 and 103). The patterns are similar for alternative definitions of "spatially moderate" MCs (see Figure S.9 in the Supplementary Materials Appendix).

These results clearly indicate that our model is not just relabeling moderate MCs as informative voters. The coexistence of partisan effects with statistically insignificant spatial differences across behavioral types may be precisely due to the fact that the ideological divisions in the House have become increasingly conflated with party labels (Noel, 2014). Hence, once we account for partisanship, any association between (un)informative voting and preference extremity essentially vanishes. For the purposes of our analysis, though, the central conclusion emerging from Figures 1 and 2 is that, even when partisanship and individual biases affect voting decisions, these factors do not trump policy-relevant information for a substantial fraction of the representatives in our sample. Thus, a model assuming that the legislative setting is entirely about (partisan, ideological) conflict would miss an important aspect of the decision-making process in the U.S. Congress.

What are, then, the other factors affecting MCs' information-responsiveness? Table 1 addresses this question, reporting the impact of institutional and electoral factors on the probability that representatives condition their decisions on the quality of legislation. The upper panel of the table focuses on the role of the institutions organizing collective action in the House. In line with the expectations outlined in Section 3, majority leaders are about 7 percentage points less likely to vote according to their private signals - and more predisposed to support proposals unconditionally - than rank and file members. The leadership of the majority party supports proposals unconditionally 96% of the time, compared to 89% for the rank and file. In the minority party, in contrast, we find no significant differences between the leadership and rank and file members in terms of their probability of being assigned to behavioral type I.

Committee members are also less prone to vote informatively than representatives not involved in drafting legislation, although there is considerable variation across issue areas. For instance, the probability that members of the Judiciary Committee decide based on their private evaluation of the quality of proposals in that area is significantly higher than for non-members, and the same tends to be true for bills from the Education and the Workforce, Defense and the Health (sub)committees (see Figure S.10 in the Supplementary Materials Appendix).[16] On average, though, the estimates for leadership

[16]These results do not change if we include Bonica (2013)'s CF scores as predictors in equation 3 to account for the possibility that committee members might be preference

and committee membership imply that representatives with gate-keeping and agenda setting power have less incentives to use policy information in their decisions than other House members.

Table 1 here

As for the impact of seniority on voting strategies and outcomes, our estimates indicate that members of the majority party become slightly more inclined to use information in their decisions as they accumulate legislative experience: increasing by one standard deviation (about four congressional terms) the number of periods they served for is correlated with a 0.82% rise in $P(\theta = I)$. For representatives in the opposition, instead, a one standard deviation increase in seniority is associated with a 2.75% decline in their probability of voting informatively. In fact, minority party MCs grow 0.66% more likely to oppose proposals independently of their quality with each additional term they serve. This suggests that the "liberating" effect of seniority discussed in Section 3 is present for the majority – when resources are more important – but not for the minority party, when they are not. Altogether, since the substantial drop in information-responsiveness among more seasoned members of the opposition is not countered by a comparable increment among long-serving majority party MCs, the net effect of an increase in seniority is a reduction in the prevalence of informative voting. On average, the mean posterior probability of being assigned to behavioral type I is more than 7 percentage points higher for a newly elected member of the House than for an incumbent with 15 or more terms of legislative experience.

Representatives' electoral environment also sways whether and to what extent they incorporate policy-specific information in their voting decisions, as seen in the bottom part of Table 1. Consistent with our expectations, minority MCs in safer districts are less

outliers conditional on party affiliation.

inclined to follow their private signals than those elected in more competitive races: a one standard deviation increase in the margin of victory is associated with a 7.25 percentage point decrease in their probability of voting informatively. At the same time, the probability that the average member of the minority consistently opposes proposals regardless of their merit would rise by 0.4% for each additional percentage gain in her vote vis-à-vis the closest challenger. Although the size of the electoral win does not significantly affect the behavior of legislators in the majority, they do become more likely to rely on policy-relevant information as constituents' familiarity with their policy stances increases, in line with Canes-Wrone and Shotts (2007)'s theoretical prediction. This impact is relatively small in magnitude, though: each percentage increase in the proportion of electors uninformed about the ideology of the incumbent reduces the probability that she takes into account the merits of the proposals by 0.02%.

Our results for seniority and margin of victory imply that the combination of uncompetitive elections and low turnover in the House has a deleterious effect on quality-based voting. This is emphasized in the left panel of Figure 3, which displays the mean posterior probability of informative voting under two counter-factual scenarios. Under Scenario 1, the House is assumed to be exclusively composed of freshman MCs elected in closely fought races, while in Scenario 2 the lower chamber comprises only incumbents serving for 15 or more terms who ran unopposed in the previous electoral cycle. The figure reveals that, holding all the other variables constant, the average probability that proposals are evaluated on their merits in the second scenario (25.5%) is almost 20 percentage points lower than in the first one. To the extent that - as contended by Iaryczower, Katz and Saiegh (2013) - members of the Senate can use the information contained in House votes to shape and improve their own decisions, a configuration of the House such as the one depicted in Scenario 2 would be decidedly detrimental for the enactment of high quality legislation by Congress.

Figure 3 here

For comparison, the right panel of Figure 3 shows that despite these differences in informative voting, MCs' ideal points do not vary significantly between the two scenarios. If anything, freshman MCs elected in more competitive races tend to be more ideologically extreme on average than senior incumbents running unopposed (see also Table S.5 in the Supplementary Materials Appendix). This again highlights that information-responsiveness in our model is not equivalent to spatial or ideological moderation.[17]

Table 2 turns attention to the bill-specific covariates assumed to affect the behavior of type-I representatives through their influence on p and q. The estimates in Column 1 show that, in line with our expectations, presidential support for a policy proposal provides a strong signal of quality. House members consider bills that are publicly backed by the President about 9.5% more likely to be appropriate for the given state of the environment than those about which he adopts no explicit position, and almost 20% more likely to be of high quality than proposals opposed by the executive. The contrast between the strong positive estimate for presidential support and the insignificant marginal effect of key votes underscores that presidential position-taking does not simply raise the salience or prominence of particular initiatives (Canes-Wrone, 2001), but also conveys relevant information about the value and likely success of the proposals.

Table 2 here

Members' beliefs about the quality of legislation are also strongly and positively correlated with the proportion of minority cosponsors. A one standard-deviation increase

[17]A comparison between Table 1 and Table S.5 uncovers in fact several differences between the determinants of informative voting and the factors driving preference moderation/extremity.

in this variable is associated with a rise of almost 10 percentage points in the likelihood that the average MC believes that the bill is of high quality. Furthermore, across-the-aisle cosponsoring activity seems to be a better predictor of legislation quality than the raw number of cosponsors.

This finding raises the possibility that we might be automatically classifying compromise bills backed by both majority and minority MCs as more likely to receive the support of informative voters - and thus, to be of "higher quality" - than proposals with a strict partisan vote. This would of course be problematic, as there is no substantive reason to assume that bipartisan bills embody better policies than proposals appealing more strongly to the majority party.[18] Nevertheless, Figure 4 shows that there is no statistically significant difference in the posterior distribution of p_t or in the proportion of informative votes across the two groups of initiatives. Furthermore, as illustrated in the lower panel of the figure, this finding is robust to alternative definitions of bipartisan bills. The same conclusion emerges from Figures S.11-S.13 in the Supplementary Materials Appendix, which compare the quality of bills appealing more strongly to the majority and minority parties and display results obtained using other measures of informative support and proposal moderateness.

Figure 4 here

Table 2 (Column 2) also shows that none of the covariates in $\mathbf{w}_t$ is associated with changes in the precision of the private signals. This is not entirely surprising, since these predictors are rather coarse proxies for the complexity and information content of the bills. We do, however, find considerable variations in signal precision across Congresses and issue areas (Figure 5). Nonetheless, q averages 0.86 across all congressional sessions and policy domains, indicating that the private information about the quality of

[18]We thank an anonymous reviewer for pointing out this issue.

legislation dispersed among members of the House is quite important in practice.

Figure 5 here

5.2 Passage in the Senate

Our analytical framework posits that bicameralism is an essential reason why we can expect to observe a substantial degree of informative voting in Congress in the first place. Here we expand the analysis in Iaryczower, Katz and Saiegh (2013) by showing that the probability that a bill originating in the lower chamber passes in the Senate depends on its support among informative House members, on these representatives' judgments about the quality of the proposal and, ultimately, on the covariates affecting MCs' beliefs and equilibrium behavior.

In order to do so, we start by estimating a simple logit model where the dependent variable $\tilde{y}_t \in \{0, 1\}$ - indicating the Pass/Fail outcome in the Senate for bills that cleared the lower chamber - is regressed against p_t and the net informative House tally τ_t, calculated as the number of "yea" minus "nay" votes among type-I MCs: $\tau_t \equiv \sum_{i:\theta_i=I} y_{i,t}$.[19] The results, reported in the upper panel of Figure 6, show that the likelihood that a proposal introduced in the House is approved in the Senate does indeed increase significantly if it receives the support of a large majority of representatives voting the bill on its merits. For each percentage point increase in p_t, the probability that the bill passes in the Senate augments by 0.21%, while each unit increase in τ_t is associated with a 0.17% rise in $P(\tilde{y}_t = 1)$.

Figure 6 here

Moreover, our model postulates that while informative votes in the House provide a public signal about legislation quality for members of the receiving chamber, the decisions

[19] The model also includes Congress- and issue-area random intercepts to account for unobserved heterogeneity in passage rates.

of non-informative MCs do not contain information that can help improve senators' choices. Hence, outcomes in the Senate should not be responsive to the tally of non-informative House votes. Consistent with this argument, the lower-left panel of the figure shows that the net support for proposals among type-Y and type-N representatives has no significant impact on $P(\tilde{y}_t = 1)$. In consequence, the effect of each additional unit increase in τ on the probability that a bill clears the Senate is significantly larger - more than twice as large - than the marginal effect of an increase in the overall House tally.

The lower-right panel of Figure 6 also shows that the support of spatially centrist House members - those with DW-NOMINATE scores below 0.25 in absolute value - has no significant effect on $P(\tilde{y}_t = 1)$, further stressing the distinction between informative voting and preference moderateness found in Section 5.1.[20] More generally, a purely ideological or spatial account of Congressional decision-making cannot explain the correlation between voting outcomes in the House and the Senate that is at the core of our theoretical model and that, as we have demonstrated, is backed by the data. In this direction, Figure S.15 in the Supplementary Materials Appendix compares the actual passage rates of House bills in the Senate against our predictions and the predictions from the spatial voting model. The figure shows that the latter does a very poor job at explaining observed data patterns, with prediction errors about 45% larger on average than for our model.

Finally, we use posterior predictive simulations (Gelman and Hill, 2007) to compute the expected change in $P(\tilde{y}_t = 1)$ associated with a change in the covariates affecting House members' responsiveness to policy-specific information. The results of this exercise are summarized in Figure 7.

Figure 7 here

[20]The same result holds for alternative definitions of moderate MCs as well (see Figure S.14 in the Supplementary Materials Appendix).

In line with the findings presented above, representatives' institutional position and electoral environment have a significant influence on the probability that House bills are approved in the Senate. Consider again a change in the composition of the House, from a situation in which all MCs served for at least 15 periods and ran unopposed in their districts, to one in which all MCs are freshmen elected in races decided by less than 1% of the vote. Everything else equal, this change is associated with a 7 percentage point boost in $P(\tilde{y}_t = 1)$. The impact of some of the bill-specific covariates is quite substantive as well. For example, the average probability that the proposal is approved by the Senate is 8 percentage points higher if the majority of cosponsors in the lower chamber belong to the minority, compared to a scenario in which only majority party members cosponsor the initiative.[21] Similarly, presidential support for the bill is associated with a 4.5 point increase in $P(\tilde{y}_t = 1)$.

6 Conclusion

Although scholars have long established that bringing about good public policy is one of the main goals pursued by members of Congress, virtually all the empirical literature in this area has focused on purely ideological or distributional problems, disregarding the quality dimension of legislation. In this paper we contribute to fill this gap, structurally estimating a model of voting that accounts for uncertainty and private information about the quality of the proposals receiving a roll call vote in the U.S. House.

Our results indicate that a sizable fraction of members of Congress evaluate proposals on

[21] We must note, though, that the p-value of a test of independence between the partisan nature of House bills and their fate in the Senate is 0.38. This reinforces our conclusion that the distinction between high/low quality legislation is not just picking up whether proposals had a strictly partisan support in the originating chamber.

their merits at the moment of casting a vote, and that their propensity to do so is fundamentally influenced by their institutional position and their electoral context, as well as by characteristics of the initiatives under consideration. In particular, agenda setting and gate-keeping power are negatively related to the probability of incorporating quality considerations in voting decisions, while competitive elections and a more junior composition of the House substantially increase the role of information in policy-making. This latter finding provides a rationale in favor of reforms aimed at increasing actual and potential renewal of the membership as a way of improving public policy. We also show that the perceived quality of House bills and the proportion of legislators following their private signals convey relevant information about the proposals for members of the Senate, and correlate with the likelihood that these bills are approved in the receiving chamber.

Much work remains ahead to better understand the intricate relations between the environment in which legislators operate, their voting behavior and policy outcomes. From a theoretical perspective, a limitation of our approach is that it does not incorporate heterogeneity in the precision of signals across legislators, a clearly restrictive assumption. Accounting for heterogeneous signals is a desirable yet challenging potential development, since characterizing the equilibrium of the model becomes substantially more complex. From an empirical standpoint, it would be relevant to include in our analysis bills initiated in the Senate in order to compare the prevalence and determinants of information-based decision-making in both chambers. Given that the number of roll call votes on passage is noticeably smaller in the upper chamber, this would require extending the period covered in our study. Another promising application of our model would be to the study of informative voting in legislatures characterized by different sets of institutions, rules and partisan compositions, contributing to our understanding of the role of information in congressional politics from a comparative perspective.

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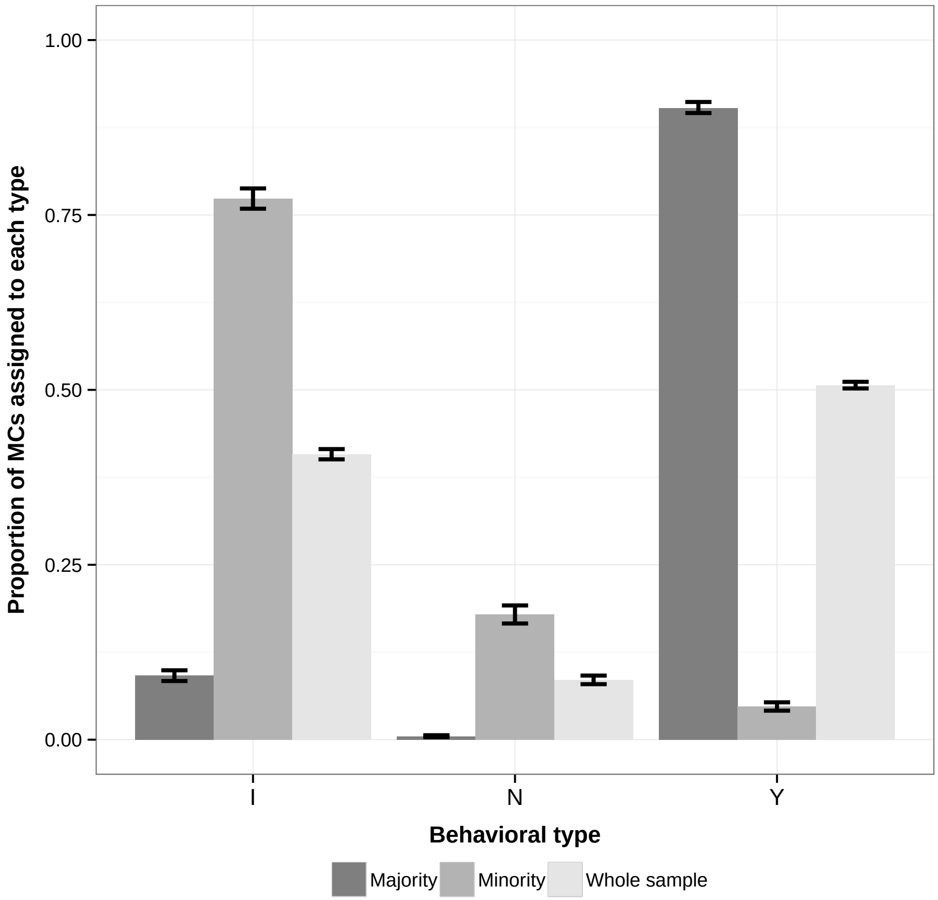

Figure 1: Posterior distribution of behavioral types in the majority and the minority. Bars represent the prevalence of each behavioral type in the majority, minority, and the whole sample. Error bars give the 90% credible intervals representing the uncertainty in legislators' assignment into types.

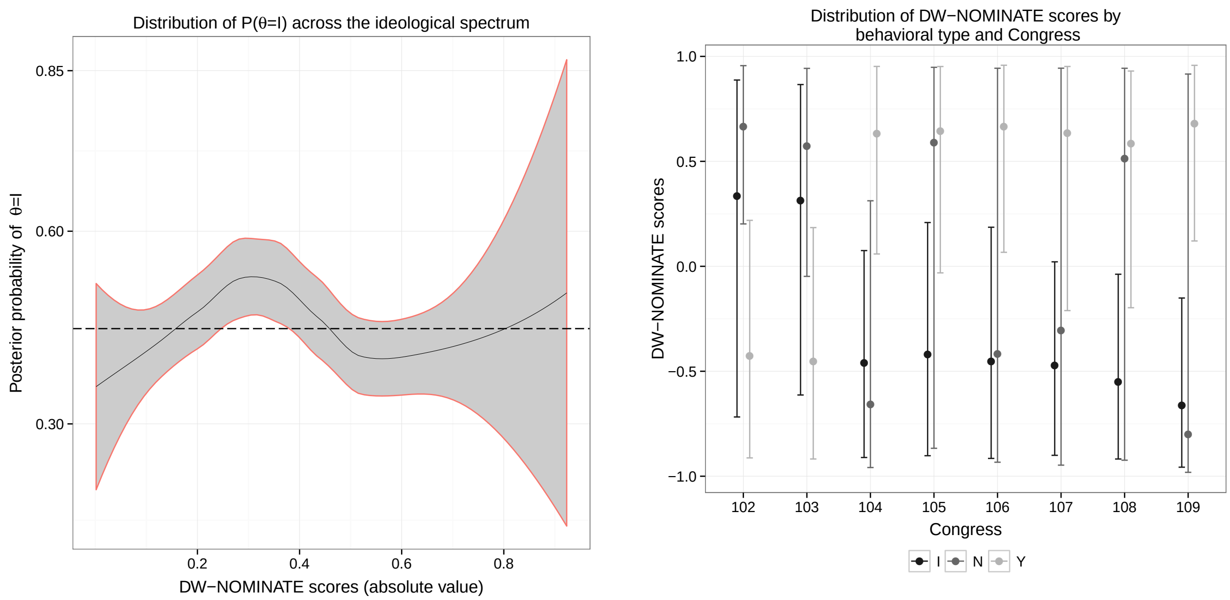

Figure 2: Relationship between MCs' behavioral types and spatial location. The left panel of the figure plots the relationship between representatives' first-dimension DW-NOMINATE scores (in absolute value) and their posterior probability of being classified as informative voters. The solid line represents the fit of a locally weighted regression curve, with 90% credible intervals given by the shaded area; the dashed horizontal line gives the average value of $P(\theta = I)$ in the sample. The right panel compares the distribution of first-dimension DW-NOMINATE scores among MCs assigned to I, N and Y types based on their maximum a posteriori probabilities of class membership. Circles represent mean DW-NOMINATE scores by behavioral type and Congress, while vertical lines correspond to the 90% credible intervals.

Table 1: Average predictive differences in $Pr(\theta_i = I)$ associated with changes in institutional and electoral variables

Covariate	(1) Majority Party	(2) Minority	(3) Whole sample
Institutional variables			
Leadership	−7.32	3.39	−2.36
	(−8.81, −5.55)	(−3.46, 9.41)	(−5.89, 0.79)
Committee Membership	−1.71	−6.21	−4.49
	(−3.49, −0.01)	(−8.29, −3.92)	(−6.18, −2.57)
Seniority	0.82	−2.75	−0.83
	(0.11, 1.53)	(−3.97, −1.52)	(−1.47, −0.13)
Electoral variables			
Margin of Victory	0.85	−7.25	−2.90
	(−0.25, 1.95)	(−8.78, −5.85)	(−3.84, −1.99)
Lack of Familiarity with the Incumbent	−0.75	0.36	−0.23
	(−1.49, −0.01)	(−0.80, 1.59)	(−0.91, 0.50)
Constituents' Level of Political Information	0.14	−1.14	−0.45
	(−0.61, 0.90)	(−2.29, 0.12)	(−1.10, 0.25)

Note: The table reports the expected percentage change in $Pr(\theta_i = I)$ associated with a change in the covariates measuring MCs' institutional position and electoral environment. Estimates (posterior means) correspond to a change from 0 to 1 in the binary covariates and to a one standard deviation increase in the continuous variables. 90% credible intervals are reported in parentheses.

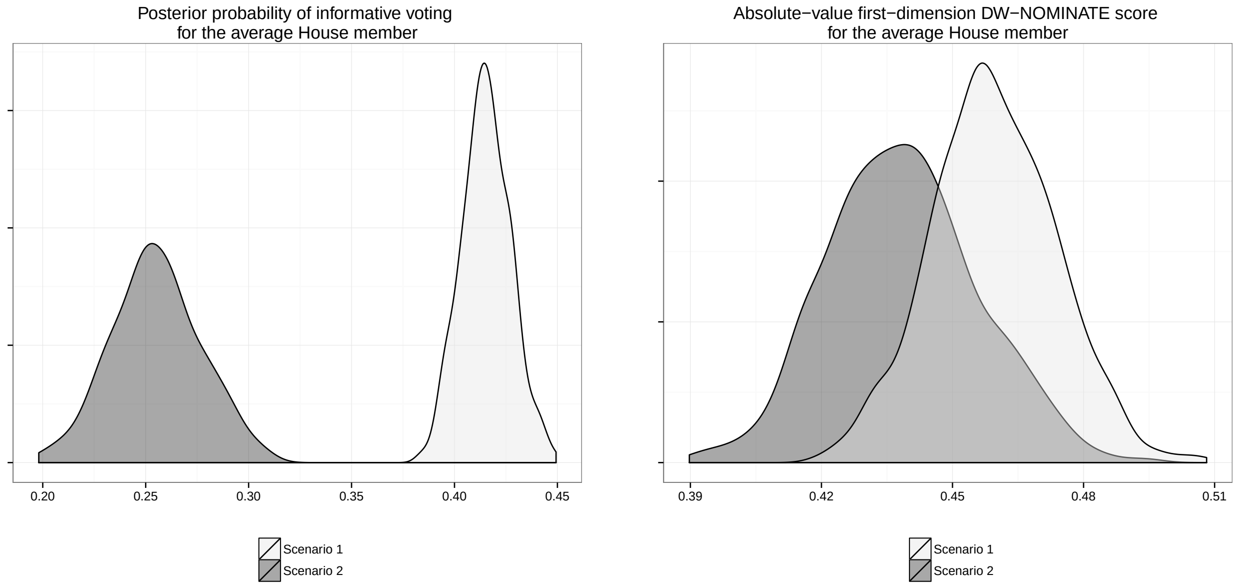

Figure 3: Informative voting and ideological preferences under two alternative compositions of the House. The left panel compares the posterior probability that the mean House member votes informatively under two counterfactual scenarios: one in which the House is composed of freshman MCs elected in races decided by less than 1% of the popular vote (Scenario 1), and one in which the lower chamber comprises only incumbents serving for at least 15 terms and who ran unopposed in the previous electoral cycle (Scenario 2). The right panel compares the ideal points of the average member of the House under both scenarios; these estimates are obtained by regressing (the absolute value of) MCs' first-dimension DW-NOMINATE scores on the determinants of θ, Congress- and constituency-level random effects.

Table 2: Average predictive differences in p_t and q_t associated with changes in bill-specific characteristics

Covariate	(1) Δp_t	(2) Δq_t
Presidential Position	9.45 (4.71, 13.54)	
Key Vote	-5.21 (-12.73, 2.31)	
Number of Cosponsors	-0.71 (-3.92, 1.98)	
% of Minority Cosponsors	9.88 (6.86, 12.79)	
Number of Words		-0.01 $(-0.12, 0.11)$
Multiple Committees		-0.01 $(-0.23, 0.21)$
Issue Experience		0.06 $(-0.99, 1.11)$

Note: The table reports the expected percentage change in p_t and q_t associated with a change in the bill-specific covariates. Estimates (posterior means) correspond to a one unit increase in the categorical predictors and to a one standard deviation increase in the continuous variables. 90% credible intervals are reported in parentheses.

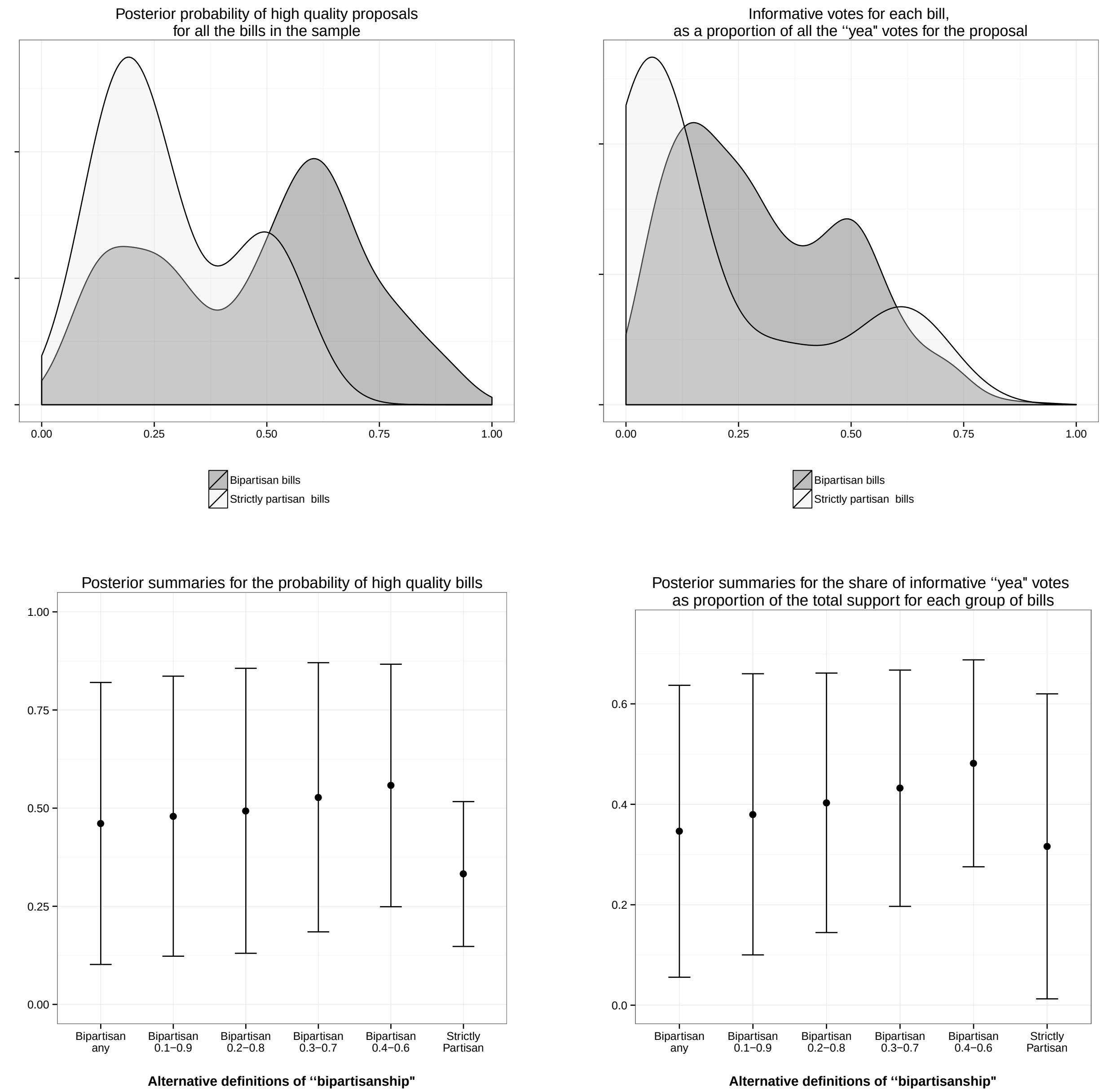

Figure 4: Proposal quality and informative support, discriminated by bills' partisan nature. The upper panel displays the distribution of p_t (left) and of the proportion of informative "yea" votes (right) for bipartisan and strictly partisan bills. The lower panel examines the sensitivity of these relationships to alternative definitions of bipartisanship resulting from changes in the proportion of majority and minority MCs supporting the bill. "Bipartisanship any" includes all bills receiving "yea" votes from members of both the majority and the opposition; "Bipartisanship x-y" corresponds to proposals receiving the support of no less than $x\%$ and no more than $y\%$ of each party's membership, with $x \in [10-40]$ and $y \in [90-60]$. Circles represent posterior means; vertical lines give the 90% credible intervals.

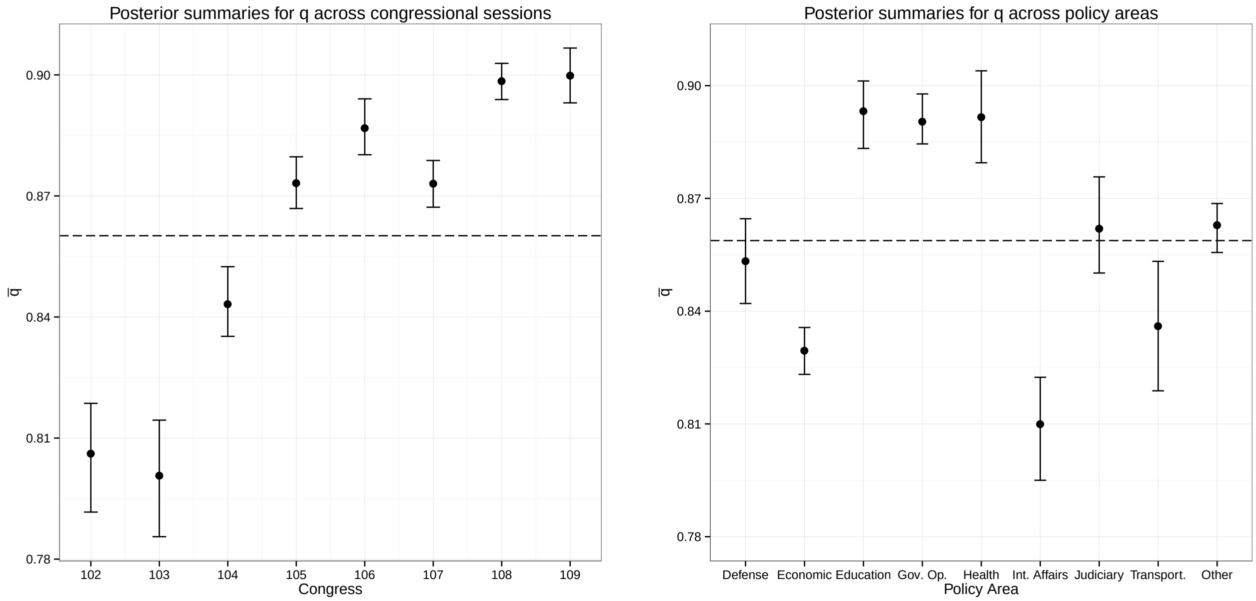

Figure 5: Posterior summaries for the precision of MCs' private information. Solid circles represent the posterior mean of q in each congressional session (left panel) and policy area (right panel). Vertical lines represent the 90% credible intervals. Dashed horizontal lines give the posterior mean of q across all Congresses and areas.

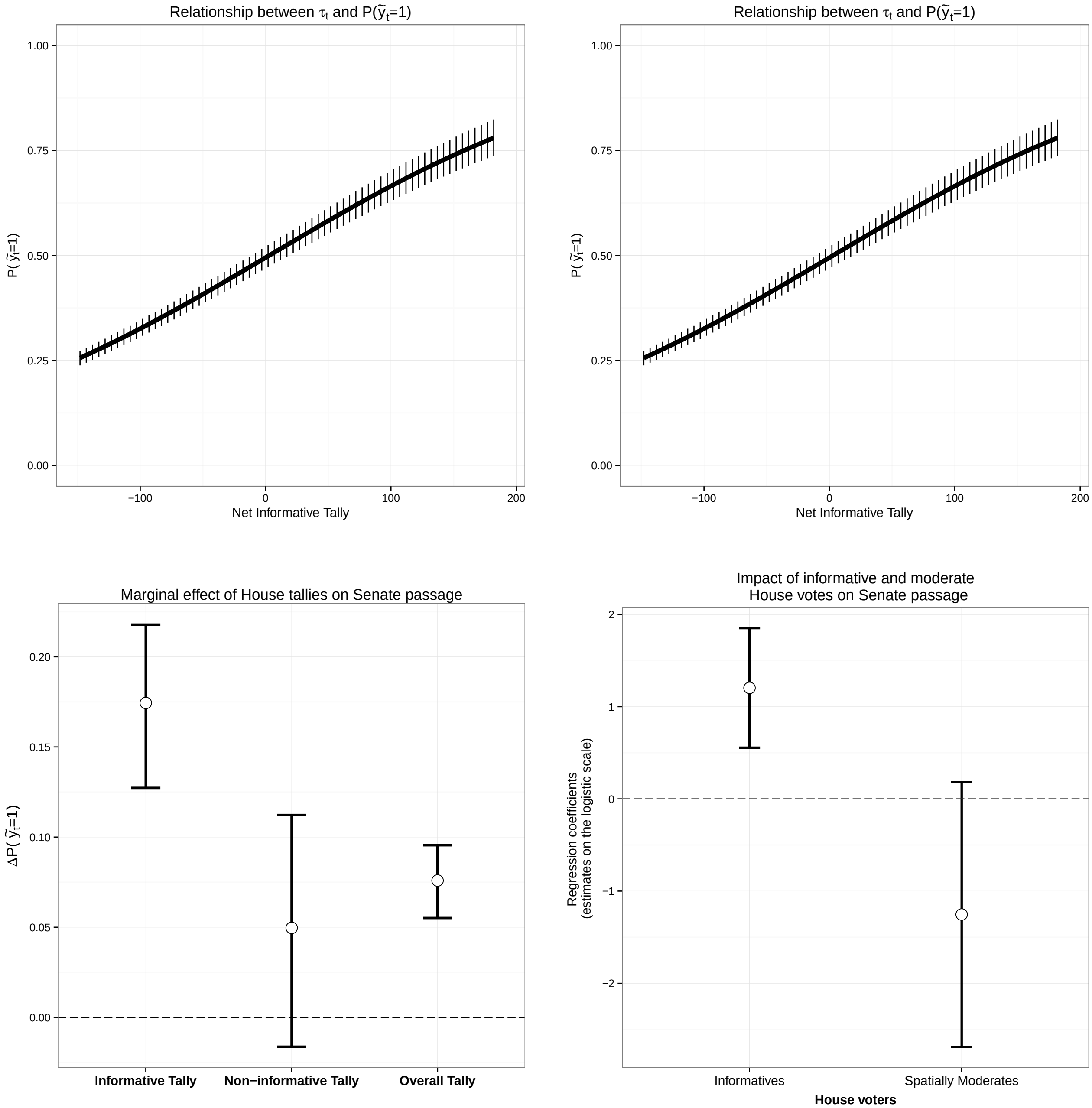

Figure 6: Impact of House bills' quality and support on Senate passage. The upper panel plots the relationship between p_t (left), τ_t (right), and $P(\tilde{y}_t = 1)$. Thick horizontal lines represent point estimates (posterior means), and the thin vertical lines correspond to the 90% credible intervals. In the lower panel, the left plot compares the expected percentage change in $P(\tilde{y}_t = 1)$ associated with a unit increase in τ_t, in the net tally among non-informative voters, and in the overall House tally. The right plot compares the coefficients for the proportion of informative and spatially moderate House votes in favor of the proposal obtained from hierarchical logit regression models for $\tilde{y}$. Circles correspond to point estimates, and vertical lines give the 90% credible intervals.

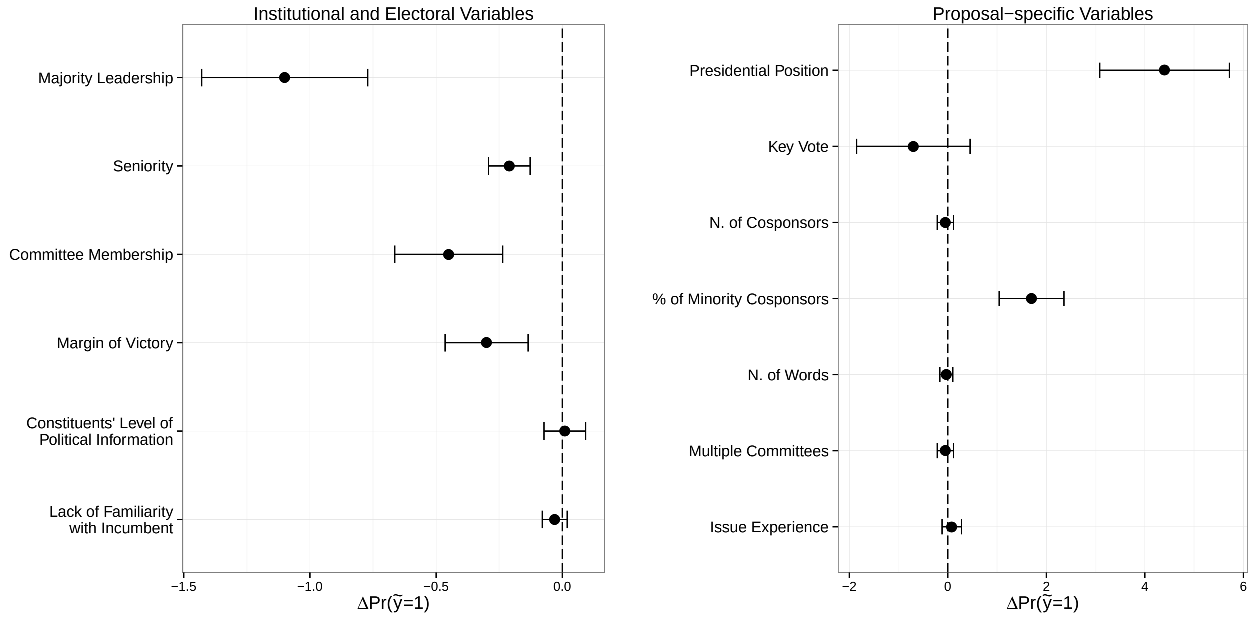

Figure 7: Impact of institutional, electoral and bill-specific variables on the probability that a House bill is approved in the Senate. Solid circles represent the expected percentage change in $P(\tilde{y}_t = 1)$ associated with a change in the covariates (a one unit increase in the categorical variables, a one standard deviation increase for continuous predictors). Horizontal lines give the 90% credible intervals.