

The Dynamics of Cosponsorship Reconsidered

by

Lawrence S. Rothenberg
Kellogg School of Management
Department of Management & Strategy
Northwestern University
lawrence-rothenberg@kellogg.northwestern.edu

and

Mitchell S. Sanders
Department of Political Science
and
Laboratory for Social Research
University of Notre Dame
msander1@nd.edu

Prepared for delivery at the 2002 Annual Meeting of the American Political Science Association, August 29 – September 1, 2002. Thanks to Christian Grose for research assistance.

Abstract

Scholars have found that specifying how legislator objectives, even if there is agreement on what they are, translate into observed behaviors is frequently problematic. Indicative, how goals such as procuring reelection are reflected in member cosponsorship choices is widely disputed. Although a fair amount of work specifying the covariates of cosponsorship exists, there is no consensus regarding inferences to be drawn from such results. Essentially two views of cosponsorship—*matching* and *signaling*—have emerged, each with a different theoretical foundation and positing a different relationship between electoral goals and behavior. Given matching, externally-oriented, reelection-minded legislators, support bills close to their ideal points as conditioned by member and electoral factors influencing their costs and benefits of participating. As such, matching defines cosponsorship as a function of the ideological or partisan match between member and bill and the costs and benefits of the participatory act of sponsoring or cosponsoring. With signaling, internally-oriented, but still reelection-motivated, members transmit information to fellow legislators via cosponsorship; as a result, extremists are more likely to cosponsor early in a bill's evolution, as their actions provide more informational content than those of moderates, and extremist-moderate differences then diminish over the period in which the bill is considered.

Our analysis takes another look at this debate by contrasting the matching and the signaling hypotheses with temporal data from the 106th Congress (1999-2000) similar to those employed by Kessler and Krehbiel (1996) but using a substantially improved research design that, among other things, places the proposal in the policy space rather than assuming that it lies at the median. We find strong support for the matching hypothesis, with members ideologically predisposed to the proposal in question far more likely to support the bill than others and extremists, if anything, more likely to jump on the bandwagon and cosponsor later in the legislative process rather than early on. As such, our results suggest that, for cosponsorship at least, member incentives result in largely externally- rather than internally-oriented behavior.

The Dynamics of Cosponsorship Reconsidered

Legislators lead complicated lives with myriad obligations and considerations. For example, while representatives may be primarily reelection oriented, it is unclear whether certain of their actions, particularly those that constituents might find difficult to observe directly, are motivated by overtly electoral or by institutional considerations. Put differently, specifying how legislator objectives, even if we agree on what they are, translate into observed behaviors is frequently problematic.

Cosponsorship is an excellent example of a choice where the relationship between goals and actions is questioned. Although a fair amount of work specifies the covariates of cosponsorship, there is no consensus regarding what inferences should be drawn from such results.

For our purposes, we may distinguish between two views—*matching* and *signaling*—each positing a different relationship between reelection desires and behavior. Analogous to standard random utility models of roll call voting that are essentially decision theoretic (e.g., Poole and Rosenthal 1998; Rothenberg and Sanders 1999a), the key prediction from the matching perspective is that cosponsorship is a function of the ideological or the partisan match between member and bill and of the costs and benefits of the participatory act (e.g., Campbell 1982, Wilson and Young 1997, Balla and Nemachek 2000). Consequently, externally-oriented, reelection-minded legislators, will support bills close to their ideal points as conditioned by member and electoral factors influencing their costs and benefits of participating.

By contrast, the signaling perspective has roots in game theoretic models of signaling with imperfect information, particularly as applied to legislatures (e.g., Crawford and Sobel 1982; Gilligan and Krehbiel 1989).¹ Thus, scholars such as Kessler and Krehbiel (1996; see also Krehbiel 1995, Gilligan and Krehbiel 1997, Wawro 2000) maintain that, rather than position-taking for external consumption, cosponsorship choices are designed to transmit information to fellow legislators. The most notable predictions stemming from this perspective are temporal: Extremists are more likely to cosponsor early in a bill's evolution, as their actions provide more informational content than those of moderates, and extremist-moderate differences diminish over the period in which the bill is considered (the matching perspective implying no such relationship).

Our analysis takes another look at this debate by contrasting the matching and the signaling hypotheses with temporal data, similar to Kessler and Krehbiel's, from the 106th Congress (1999-2000). In doing so, we suggest that Kessler and Krehbiel's test, which we explicitly build upon, can be improved in terms of measurement, specification, and estimation. Regarding measurement, we make several changes; most dramatically, we directly measure where bills fit in the policy space rather than rely upon the implicit assumption that the cosponsored proposal is located at the legislative median. Concerning model specification, we incorporate several additional factors that existing analyses of legislative decision-making

¹Some political scientists discussing cosponsorship employ the term signaling in a more encompassing manner than ours. For conceptual clarity, we adopt a narrower usage corresponding more closely to signaling's formal game theoretic definition. For instance, while Wilson and Young (1997) place three alternative determinants of cosponsorship under the signaling rubric—bandwagon, ideological, and expertise—for our purposes, bandwagon and ideological determinants are subsumed by matching (we will discuss bandwagon effects in our empirical analysis) while only expertise is captured by signaling.

suggest are potentially important for determining member behavior—committee memberships, retirement status, and partisanship among them. And with respect to estimation, we employ a more flexible method allowing us to examine the determinants of cosponsorship in a continuous fashion, without restrictive assumptions, over a bill’s lifespan. Taken together, our analysis allows us to draw inferences confidently about whether matching or signaling is at the heart of the cosponsorship process.

When we do, indeed, examine cosponsorship and its timing we find strong evidence that legislators engage in external matching rather than in internal signaling. Although some determinants of cosponsorship vary with time, we find that, everything being equal, those whose ideal point roughly corresponds to that of the proposal in question dominate cosponsorship activity. This is especially true early in the legislative process. Later, as momentum grows, there is somewhat of a bandwagon effect by which some of those less naturally supportive, particularly liberals, come on board, presumably largely to curry constituent favor.

Our analysis proceeds in three steps. First, we briefly overview the matching and signaling perspectives, particularly by reviewing the aforementioned work of Kessler and Krehbiel whose temporal dimension is unique and corresponds to our analysis, and expound on how hypotheses associated with these two views may be more directly tested. Next, we specify an empirical model to test these alternative perspectives. Finally, we estimate hazard rate models of cosponsorship and draw inferences regarding their implications for which view best characterizes the underlying dynamic process.

Match or Signal?

To reiterate, Kessler and Krehbiel most clearly highlight the distinction between matching and signaling, maintaining that signaling predicts that bills with broad support should more likely have early cosponsorship from extremists and that matching implies no such relationship. Empirically, they assemble data on the 51 bills in the 103rd Congress (1993-1994) with at least 50 cosponsors and score whether a legislator cosponsored the bill in one of three temporal periods.² Each member-bill combination is categorized according to whether there was (0) no action; (1) early period action; (2) intermediate period action; or (3) late action, with periods determined partially by the length of time between the bill's introduction and either its passage or the end of the Congress. These 51 bills are further broken up into those passing (which should more accurately reflect the specified pattern) and binding bills having the force of law (i.e., excluding resolutions, which may be less relevant because they lack the force of law if approved). Members, in turn, are classified according to their ideological extremity (by quartile, using ADA scores), seniority (in years), and electoral margin (natural log of percentage). With this specification, two semi-parametric hazard rate models are then run—an unconstrained model allowing for different behavior by liberals and conservatives, on the grounds that liberals may have more activist dispositions, and a constrained model imposing symmetry—and it is found that extremists are both more likely to cosponsor and to do it early, with effects more forceful for liberals. Thus, Kessler and Krehbiel conclude that “legislators do not use bill sponsorship as a

²Although it is understandable why one might select bills with many cosponsors, such a selection criteria may also be unrepresentative. Therefore, we will start with the universe of proposals and then check the robustness of results by examining proposals with high cosponsorship levels.

mechanism for position taking aimed primarily at external audiences” (p. 563) but rather adopt it as an internal signaling device.

Although an innovative analysis, as implied, there are reasons to take another look at cosponsorship dynamics. While there are various elements of the research designed that can be improved, the most notable involves proxying the position of the proposal in the policy space. Put differently, are conservatives merely supporting conservative bills, liberals supporting liberal bills, and moderates supporting moderate bills? Kessler and Krehbiel’s claim that extremists should jump in first and that moderates should come in later is predicated on implicitly assuming that the initial bill is at the median. Alternatively, it makes more sense to assume that, even with the constraints posed by the need to generate support for eventual passage, relatively extreme members actually propose relatively extreme bills, some of which are modified as time goes on in anticipation or hope of eventual passage.³ This implies that if extremists on the side of the fence consistent with where a bill is proposed jump in early, with perhaps some moderation toward the center as time elapses, some moderation is taking place.⁴ While Kessler and Krehbiel try and formulate some indirect tests of these propositions, they do not directly integrate information on where a bill fits in the policy space.

³Content analysis suggests that there are some proposals in our sample which are modified toward the center but, even for initiatives with many cosponsors, this is a small percentage.

⁴Ideally, we would know the status quo policy as well so that for any member we could estimate a Euclidean distance comparing her ideal point, the policy proposal, and the status quo. Unfortunately, locating the status quo is extremely difficult so we will have to assume that, everything being equal, legislators will support proposals closer to their own ideal points.

To proxy the policy position at which a bill is located, we employ the proposer's ideological position (e.g., Pelligrini and Grant 1999).⁵ Assuming that a sponsor's ideology reflects where the introduced bill falls on the policy space allows us to incorporate the distance between a member and the initial policy proposal into our hazard rate analysis. Put differently, we define an extremist as a representative who is ideologically distant from the initial sponsor and not merely as a legislator who is extreme from the median member.

In addition, as mentioned, we improve upon past measurement and specification. As for measurement (and with specific reference to Kessler and Krehbiel), we employ continuous rather than truncated measures of ideology (both of sponsors and potential cosponsors) using superior Poole-Rosenthal W-NOMINATE scores instead of ADA measures. We also measure tenure non-linearly on the presumption that, like electoral margin (which, to reiterate, Kessler and Krehbiel measure non-linearly), there may be a diminishing impact associated with years of service.⁶ Perhaps most importantly, rather than divide a bill's lifespan into early, intermediate,

⁵Detailed content analysis suggests that, although not perfect, this is a rather good predictor of where the proposal falls on the ideological space. The biggest discrepancy appears to be for certain high profile initiatives where there is *de facto* joint sponsorship across the policy spectrum but where, according to House rules, only one legislator is the technical sponsor.

⁶A further point to mention is that, with respect to the overall impact of security and tenure on cosponsorship levels, and although not commonly recognized in the relevant literature, expectations are not clearcut given that moral hazard and selection effects should work at cross-purposes. Members may be secure because they are good representatives and, as one manifestation of their high quality, engage in much cosponsorship activity; alternatively, security may be a function of reasons exogenous to representative quality and, as such, secure members may generally work less and may specifically cosponsor less. Those with lengthy tenure may have survived the reelection process because they are good representatives and, therefore, cosponsor more; conversely, senior members, getting closer to the end of their legislative careers may cosponsor less because they have less incentive to engage in costly activity.

and late periods, we examine behavior in continuous time throughout the period the bill is under consideration as a means of getting a fuller picture of temporal dynamics.

Additionally, we incorporate a number of other obvious features related both to the propensity of members to cosponsor generally and to the utility of cosponsoring a given bill specifically into our specification. Concerning the former, for example, we include whether a representative is seeking reelection, as the literature on legislator incentives suggests that last-term members may reduce their effort, implying that endgaming can induce declining cosponsorship (e.g., Poole and Rosenthal 1997; Rothenberg and Sanders 1999b, 2001). Furthermore, we include party membership as research on political parties, albeit controversial, indicates that being in the majority or minority party might be relevant for the kinds of the behavior that we observe as the majority party is claimed to have asymmetric influence (for a review, see Aldrich and Rohde 2000; see, also, Krehbiel 1998; Binder, Lawrence, and Maltzman 1999). In a similar vein, but with respect to factors that might influence cosponsorship of specific proposals, we also include whether there is a match between the party of the proposal's sponsor and the party of the cosponsor on the grounds that members may have more to gain in supporting a fellow party member's bill.

Finally, econometrically, we improve on the means that Kessler and Krehbiel use by adopting a flexible approach. Hence, while we too employ a hazard model, we adapt it to incorporate a more complete notion of time: Where Kessler and Krehbiel divide time for each proposal into the same number of periods (the length in real time for each period being determined by the proposal's lifespan), and thus assume that hazard rates for cosponsorship are constant for long stretches of time, we measure time continuously in the common metric of days

since introduction and allow hazard rates to vary continuously as well. Also, rather than assuming that the effects of variables other than ideology are constant for all levels of ideology, we estimate separate models for various strata of member-proposal ideological match that allow both the shape of the hazard curve and the impacts of covariates to fluctuate.

Data, Measurement, and Estimation

As a first step in our analysis, we code all House sponsorship and cosponsorship activity for the 106th Congress by proposal and House member. We include all House Bills, House Concurrent Resolutions, and House Joint Resolutions introduced by virtually all [421] legislators.⁷ Of the 5,859 qualifying proposals, 90.3 percent are House Bills, 7.4 percent are House Continuing Resolutions, and 2.3 percent are House Joint Resolutions.⁸ With respect to the entire sample of proposals, the average number of cosponsors is 19.3, with a standard deviation of 38.9. Overall, 30.1 percent of proposals had no cosponsors, 1.6 percent only had cosponsors upon initial introduction, and 68.3 percent had cosponsors after initial introduction. Cosponsorships on the day of introduction accounted for 35.2 percent of the total and the average

⁷As implied, after our initial coding, we exclude some members (and the bills they sponsor): the House Speaker, those exiting before the end or arriving after the beginning of the Congress, those switching parties or who were independents during the Congress, those not voting often enough to receive W-NOMINATE scores, the Louisiana delegation (because of that state's unique electoral system), and nonvoting delegates. Also, we exclude each proposal's sponsor from the roster of potential cosponsors, leaving 420 potential cosponsors per bill.

⁸We exclude private bills and bills introduced by request as they are unlikely to be characterized by the processes of interest.

cosponsor signed on 71.4 days after a proposal's introduction (with a standard deviation of 111.1 days).⁹

Our dataset thus consists of each member-proposal combination (2,460,780 observations) for whether and when a proposal is cosponsored by members not serving as sponsor. For each member and each proposal the member can cosponsor before the ending of the eligible period for cosponsorship or can abstain. If the member cosponsors, we code the date on which that action occurs. As proposal introduction and reporting can take place throughout a Congress, the precise timing of these activities varies. To account for this, we rescale each proposal so that its introduction day is coded as day 1 and a member's cosponsorship date, *Days Since Introduction*, is coded as the number of days from when the introduction occurs.¹⁰ For example, regardless of when a proposal is introduced, we code initial cosponsors as acting on $t=1$, those cosponsoring on the second day as deciding on $t=2$, etc. Consistent with standard application of duration models, we treat observations for non-cosponsoring members as censored, with the observation period ending either when floor action is taken on the proposal (ending the opportunity to cosponsor)¹¹ or, given no action, when the 106th Congress concludes.

To reiterate, for independent variables, we measure features associated with the incentive to cosponsor generally and to cosponsor on specific kinds of proposals. By controlling for these

⁹Not surprisingly, cosponsorship's temporal distribution is skewed: The median time to cosponsorship was 23 days, 77.3 percent of all cosponsorship occurred within 100 days of introduction, and 89.3 percent occurred within 200 days.

¹⁰As will be seen shortly, we use this variable interactively with our independent variables to correct for possible violation of the proportionately assumption of hazard models.

¹¹Technically, cosponsorship opportunities end on the day on which the proposal is reported by a committee to which it was referred.

factors, our model allows us to test whether matching or signaling is taking place by investigating the relationship between proposal-specific ideological extremity and temporal cosponsorship.

Concerning the general propensity of legislators to cosponsor, we specify four variables to capture how costs and benefits vary across members:

- (1) *Electoral Security*, measured as the natural logarithm of the percentage of the two-party vote that the member received in the 1998 election;
- (2) *Seniority*, measured as the natural logarithm of the number of years of member service in the House;
- (3) *Party*, coded as one for Republicans and as zero for Democrats; and
- (4) *Lame-duck*, measured as whether, at the time of the proposal's introduction, the member planned to retire at the end of the 106th Congress.¹²

As mentioned, expectations regarding security and seniority for overall sponsorship are ambiguous, depending upon whether selection or moral hazard effects are stronger on net.

Analogously, expectations for party are ambiguous: Democrats may be thought of as more prone

¹²Measuring lame-duck status is a bit tricky for two reasons. First, as Rothenberg and Sanders (1999b) demonstrate, assuming retirees behave as lame-ducks throughout the whole two-year congressional session and that those seeking reelection were certain that they would continue introduces at least some measurement error. We avoid much of this problem by looking at whether the member had decided to retire by the date of a proposal's introduction. This raises the second issue, when to code members as retiring. To measure this, for each exiting member we conducted an intensive investigation of published sources, both of national and local newspapers and magazines, to ascertain when members had apparently made their retirement decisions and used this information to code lame-duck status. Some members, for example those who had previously announced that they would be term limited and stuck to this promise without wavering, were coded as having chosen to retire for the entire Congress while others, for example members in New Jersey and New York who had an opportunity to run for Senate when the front-running candidate unexpectedly dropped out of the race, were coded as choosing retirement much later.

to cosponsor, everything being equal (especially as we are not measuring ideology as an independent variable), but Republicans, being the majority party, might be induced to cosponsor more given their potential agenda-setting advantage.¹³ Lame-duck members should be less likely to engage in cosponsorship, particularly once we control for seniority, regardless of whether one adopts the signaling or matching perspective.

As for issue-specific factors, we incorporate two:

- (1) *Committee membership*, scored one if the member sits on at least one committee considering the proposal and zero otherwise; and
- (2) *Sponsor partisanship*, a dummy variable scored one if a member is of the same party as the initial sponsor and zero otherwise.

In both instances, expectations for the effect of these factors on overall sponsorship are positive, as committee membership indicates that a member has more at stake in a proposal than other legislators while sponsor partisanship essentially serves as a proposal-specific measure of partisan influence. We also expect that committee membership will be associated with early cosponsorship as, minimally, the member is confronted earlier in the legislative process than most members with the cosponsorship choice.

As foreshadowed, while ideological distance from the proposal's sponsor is the key given our theoretical interest, we use it to stratify our sample rather than incorporate it directly as an independent variable. To reiterate, this allows us flexibility, as the hazard rate function's shape

¹³Our results are essentially unchanged even if we drop party because it is proxying for ideology.

can now vary with ideological distance, allowing us to determine if extremists or centrists are more likely to cosponsor overall and at different points in the statutory process.¹⁴

Specifically, we incorporate ideological distance using a three-step process. First, we specify two variables, *Member Ideology* and *Sponsor Ideology*, using first dimension W-NOMINATE scores for the member and the sponsor respectively (as is well-known, these scores lie roughly in the interval [-1,1], with negative and positive values representing liberal and conservative predispositions). We then calculate *Ideological Difference* = *Member Ideology* – *Sponsor Ideology* for each member-proposal combination, giving us a measure of ideological correspondence between legislator and proposal ranging from -1.969 to +1.969, with a mean of -0.053 and a standard deviation of 0.816. Finally, we divide *Ideological Difference* into quintiles (the boundaries between the quintiles being -0.908, -0.182, 0.117, and 0.686). As should be clear from this measure’s construction and its resulting distribution, the first and fifth quintiles represent member-proposal combinations where the member is much more liberal and much more conservative than the proposal, the second and fourth involve members being somewhat more liberal and conservative, and the third reflects members having a close ideological correspondence with the proposal.¹⁵

¹⁴For example, this flexibility circumvents Kessler and Krehbiel’s concern that including preferences as an independent variable will stretch or compress baseline hazards without allowing the basic shape of these hazards to change.

¹⁵This measure results in considerably different placement of legislators relative to the bill than simply using the median. For illustration, if we simply take our 421 legislators and categorize them as liberals (the most liberal 25 percent as defined by their NOMINATE scores), moderates (the middle 50 percent on the NOMINATE scale), and conservative (the most conservative 25 percent), and look how often they fall into different quintiles, we find that liberals are in the most liberal quintile 54.9 percent of the time, the next most liberal quintile 21.7 percent of the time, the middle quintile 16.9 percent of the time, the fourth quintile 6.6 percent of

Having delineated *Ideological Difference* quintiles, we run a hazard rate model for each separately. We base our model on the semi-parametric Cox proportional hazard specification, where for each observation the hazard rate at time t is given by:

$$h(t) = h_0(t) \exp(b_1 X_1 + b_2 X_2 + \dots + b_k X_k) \quad (1).$$

That is, the hazard rate is determined by the baseline hazard $h_0(t)$ and the observation-specific effects measured by X_1, X_2, \dots, X_k . However, we deviate from the original Cox specification by allowing covariate effects to vary over time (using *Days Since Introduction*) rather than assuming that each proportional effect on the hazard is the same at all times, meaning that the hazard rate at t is:

$$h(t) = h_0(t) \exp(b_1 X_1 + b_2 X_2 + \dots + b_k X_k + b_{k+1} X_1 * t + b_{k+2} X_2 * t + \dots + b_{2k} X_k * t) \quad (2).$$

While the baseline hazard rate is not directly estimated, we can recover the survival functions after estimation—in this case, the cumulative probabilities of cosponsoring the bill through t as they vary for each stratum of *Ideological Difference*—and these functions enable us to examine the changing probabilities of cosponsorship over time and to identify any temporal/ideological patterns in cosponsorship. Thus, we are able to specify parametrically the effects of our variables on the hazard rate without committing to any particular functional form for the baseline hazard and to have different baseline hazards for each group.

More to the point, we can test the signaling and matching hypotheses with confidence.

To reiterate, evidence that, all else equal, members who are extremists relative to the proposal's position in the policy space are more likely to cosponsor early supports the signaling perspective.

the time, and the most conservative quintile 0.0 percent of the time. Analogous percentages by quintile for moderate legislators are 12.4, 28.2, 19.3, 19.9, and 20.2 and for conservatives they are 0.0, 1.6, 24.6, 33.8, and 40.0.

Failure to find this temporal pattern but discovery that legislators tend to support bills close to their ideal points as conditioned by individual- and proposal-specific factors is evidence of matching.

Results

Table 1 shows maximum likelihood estimates (MLEs) and associated z -scores generated by our proportional hazards models for our five strata which, for purposes of convenience, we label from very liberal to very conservative. These coefficients measure the change in the hazard rate accruing from a one unit change for the independent variable; coefficients less [more] than zero indicate that a variable decreases [increases] the cosponsorship hazard. As foreshadowed, all independent variables are included both directly and interactively with *Days Since Introduction* to allow their effects on the hazard rate to vary with time.¹⁶

Overall, the model does well. Strong chi-squared scores indicate that it represents a statistically significant explanation of cosponsorship decisions for all strata.

While all specified factors statistically influence the hazard rate for at least some strata of *Ideological Difference*, influences vary depending upon member extremism relative to the proposal, particularly for those factors, seniority and electoral security, that may be influenced both by moral hazard and selection effects. Seniority has positive effects for those who are very conservative or very liberal relative to the proposal, in both cases becoming smaller in magnitude over time, while it has an initial negative effect that becomes larger in magnitude over time for

¹⁶As implied, tests based on Schoenfeld residuals suggest that the proportional hazards assumption may be problematic, that a model without time-varying parameters is inappropriate, and that our more flexible specification allowing, but not requiring, variable effects over time is needed.

centrists. For liberals and conservatives, seniority does not have an initial effect, but the interaction term indicates that there is a growing negative effect over time. By contrast, electoral security is insignificant for liberals, leads to greater cosponsorship activity for centrists, but produces less cosponsorship when they are more conservative; the effects become smaller in magnitude over time for centrists and for extreme conservatives. As for partisanship, although its effect weakens over time, for all but the first stratum (for which a coefficient cannot be estimated as virtually all members included are Democrats), Democratic legislators are more likely to cosponsor than Republicans, indicating that any agenda-setting impacts are overcome by Democratic activism. Not surprisingly, the effects of lame-duck status are negative, as departing members are less likely to cosponsor, and roughly constant over time and across all strata. Interestingly, the effect of being in the same party as the bill's sponsor varies over the ideological strata, as it is insignificant for extreme liberals and conservatives, positive but declining over time for moderate liberals and centrists, and steadily positive for moderate conservatives. Finally, the effects of being on a considering committee are strongly positive but, not surprisingly, decreasing over time for all strata.

Although these results are interesting, of greater relevance for our analysis are the different shapes of the baseline probabilities of cosponsorship for each of the five quintiles of member-proposal extremism. Do extremists cosponsor more early and centrists more late or is a pattern more consistent with matching observed? Figure 1, which displays the probability that a member will cosponsor a proposal given its placement in the policy space relative to her preferences on or before time t holding all the independent variables at the corresponding

quintile's sample mean, indicates the latter.¹⁷ While cosponsorship's probability is, obviously, always low, and cumulative probabilities always increase over time by construction, centrists relative to the proposal are most likely to cosponsor over its lifespan, followed by moderate liberals and moderate conservatives and, finally, by extreme liberals and extreme conservatives.

Put differently, we can draw three inferences from Figure 1: (1) ideological matching dominates in a cumulative sense throughout; (2) cosponsorship exhibits a liberal bias; and, most importantly, (3) there is no pattern by which extremists [centrists] can at all be construed as entering early [late] but not late [early]. With respect to the latter point, for example, even after the disparity associated with initial cosponsorship on day 1 is taken into account, the cumulative increase in centrist probability of cosponsorship for the first 100 days exceeds 0.030 while it is under 0.015 even for extreme liberals.

As such, extremists do not dominate in either a relative or an absolute sense at a bill's early stages. If anything, we find a bandwagon effect by which initial centrist support brings others whose ideal points are further from where the proposal is in the policy space to come on board. To illustrate, we calculate the expected composition of a bill's cosponsors over time with respect to distance from the bill (for simplicity, ignoring the differences between liberals and conservatives). Put differently, for each quintile at each point in time we calculate the expected proportion of a bill's cosponsors, e.g., the expected probability of cosponsoring at $t=1$ is 0.0128 for extremists (those in the first and fifth quintiles), 0.0339 for moderates (members in the second and fourth quintiles), and 0.0265 for centrists (members in the third quintile). As such, at

¹⁷In all figures we show the entire range of a proposal's possible lifespan, but we should reiterate that most cosponsorship occurs within three to six months of introduction.

$t=1$ extremists constitute 17.5 percent of cosponsors, centrists 36.2 percent, and moderates 46.3 percent (remember that centrists are only 20 percent of our sample by this construction, compared to 40 percent for extremists and moderates).

As Figure 2 illustrates, the evidence is inconsistent with the signaling perspective but is consistent with matching where there is a bandwagon effect by which members not obviously ideologically disposed are more prone to get on board as support grows. Specifically, the proportion of centrist cosponsors decreases over time, the proportion of moderates grows slightly, and the proportion of extremists increases.¹⁸ To reiterate, if, as the signaling perspective suggests, extremists are early cosponsors, then we would expect exactly the opposite: the proportion of extremists would be highest and the proportion of centrists would be lowest early in a proposal's lifespan, then the former would decrease and the latter would increase as time passes. And while one might argue that this is some odd sort of reverse signaling, by which centrists signal extremists that the bill is worthwhile, this is refuted by the fact that the growth of extremist cosponsors is almost exclusively from those more liberal relative to proposals (juxtaposing the evidence in Figures 1 and 2), an asymmetry which is consistent only with a matching perspective by which legislators from liberal districts disposed toward government intervention join in when probability of enactment grows.¹⁹

¹⁸In addition, notice that extremists are underrepresented throughout a bill's lifespan—while we define extremists as composing 40 percent of our sample, they never comprise more than about 20 percent of cosponsors.

¹⁹Krehbiel and Kessler suggest that “inside-out” cosponsorship, where extremists on both sides of the ideological distribution get on board after initial centrist support, might reflect entrepreneurial behavior on the part of moderate party leaders (but not signaling). However, the liberal bias of extremist cosponsorship seems to support a more straightforward matching perspective.

Of course, our results are based on a very large and diverse sample of proposals. Might the signaling perspective receive support for certain subsets of proposals for which information transmission signaling might be crucial? As such, and corresponding to Kessler and Krehbiel's analysis, we check the robustness of our central results by restricting our analysis to the 11.2 percent of proposals with 50 or more cosponsors (these proposals encompass 61.9 percent of the total number of cosponsorships). As we can see from Figure 3 (presented analogously to Figure 2 to facilitate comparison), our results are not artifacts of taking a broad sample. The proportion of centrist cosponsors declines with time, the proportion of moderates remains nearly constant, and the proportion of extremists rises. These results provide more evidence for inferring that matching rather than signaling characterizes cosponsorship.

In sum, while ideological extremism of members relative to bills conditions the effects of other variables on the hazard rate of cosponsorship, it does so in a manner consistent with matching and not signaling. All in all, there is no evidence that internal information transferral is key for understanding cosponsorship. Rather, matching with constituencies, be they organized supporters or rank and file voters, by which ideological correspondence and other conditioning factors are crucial, appears to be of paramount importance.

Discussion and Conclusions

Understanding how basic member incentives translate into behavior has often proven problematic generally and with respect to cosponsorship specifically. Although no analysis is definitive, our results lend strong credence to the belief that such puzzles can be solved, at least

for cosponsorship. Specifically, by utilizing earlier insights but remedying research design deficiencies, we show that there is no support for internal dynamics being crucial but that there is consistent and undeniable evidence that members use cosponsorship to communicate with external constituencies. While we do not wish to suggest that internal signaling does not occur in the United States Congress, our results indicate that it is not through the mechanism of cosponsorship.

There would seem to be good, intuitive, reason why cosponsorship would not be the informational vehicle of choice. First, the cost of cosponsorship may be extremely low and therefore incorporate little informational content. Given this low cost, it may often be difficult for fellow cosponsors to distinguish between informational cosponsoring and cosponsorship undertaken for other reasons, discouraging attempts to use cosponsorship for internal communication altogether. Second, while many voters might not view cosponsorship directly, it is easily observable by electorally-relevant elites, be they interest groups or political candidates—for example, an extremist member’s cosponsorship of a proposal that proves quite unpopular to her supporters could be easily exploited by a challenger while the same member can trumpet her cosponsorship of an initiative that is consistent with district or interest group preferences.

References

- Aldrich, John, and David Rohde. 2000. "The Consequences of Party Organization in the House: The Role of the Majority and Minority Parties in Conditional Party Government," in *Polarized Politics: Congressional and the President in a Partisan Era*, eds. Jon R. Bond and Richard Fleisher. Washington, D.C.: Congressional Quarterly.
- Balla, Steven J., and Christine L. Nemachek. 2000. "Position Taking, Legislative Signaling, and Nonexpert Extremism: Cosponsorship of Managed Care Legislation in the 105th House of Representatives," *Congress & the Presidency* 27: 163-188.
- Binder, Sarah A., Eric D. Lawrence, and Forrest Maltzman. 1999. "Uncovering the Hidden Effects of Party," *Journal of Politics* 61: 815-831.
- Campbell, James E. 1982. "Cosponsorship Legislation in the U.S. Congress," *Legislative Studies Quarterly* 7: 415-422.
- Crawford, Vincent P., and Joel Sobel. 1982. "Strategic Information Transmission," *Econometrica* 50: 1431-1451.
- Gilligan, Thomas W., and Keith Krehbiel. 1989. "Asymmetric Information and Legislative Rules with a Heterogeneous Committee," *American Journal of Political Science* 33: 459-490.
- Gilligan, Thomas W., and Keith Krehbiel. 1997. "Specialization Decisions Within Committee," *Journal of Law, Economics & Organization* 13: 366-386.
- Kessler, Daniel, and Keith Krehbiel. 1996. "Dynamics of Cosponsorship," *American Political Science Review* 90: 555-566.
- Krehbiel, Keith. 1995. "Cosponsors and Wafflers from A to Z," *American Journal of Political*

- Science* 39: 906-923.
- Krehbiel, Keith. 1998. *Pivotal Politics: A Theory of U.S. Lawmaking*. Chicago, IL: University of Chicago Press.
- Pellegrini, Pasquale A., and J. Tobin Grant. 1999. "Policy Coalitions in the US Congress: A Spatial Duration Modeling Approach," *Geographical Analysis* 31: 45-66.
- Poole, Keith T., and Howard Rosenthal. 1997. *Congress: a Political-Economic History of Roll Call Voting*. Oxford: Oxford University Press.
- Rothenberg, Lawrence S., and Mitchell Sanders. 1999a. "Rational Abstention and the Congressional Vote Choice," *Economics & Politics* 11: 311-340.
- Rothenberg, Lawrence S., and Mitchell Sanders. 1999b. "Severing the Electoral Connection: Shirking in the Contemporary Congress," *American Journal of Political Science* 44: 310-319.
- Rothenberg, Lawrence S., and Mitchell Sanders. 2001. "The Political Economy of Congressional Careers." Unpublished manuscript.
- Wawro, Gregory. 2000. *Legislative Entrepreneurship in the U.S. House of Representatives*. Ann Arbor, MI: University of Michigan Press.
- Wilson, Rick K., and Cheryl D. Young. 1997. "Cosponsorship in the U.S. Congress," *Legislative Studies Quarterly* 22: 25-44.

Table 1:
Effects on Cosponsorship Hazards, by Ideological Distance from Proposal
(MLEs from Cox Proportional hazard Model, with z-scores in Parentheses)

	<i>Member Extremism Relative to Proposal (by Quintile)</i>				
<i>Variable</i>	Very Liberal	Liberal	Centrist	Conservative	Very Conservative
log(Seniority)	0.036* (3.46)	0.013 (1.58)	-0.025* (-3.55)	-0.010 (-1.09)	0.16* (8.92)
log(Seniority)* Days Since Introduction	-0.00044* (-6.38)	-0.00027* (-4.58)	-0.00015* (-2.58)	-0.00039* (-5.94)	-0.00057* (-4.34)
log(Electoral Security)	0.037 (0.74)	0.0014 (0.04)	0.090* (2.94)	-0.13* (-3.58)	-0.63* (-8.54)
log(Electoral Security)* Days Since Introduction	0.00014 (0.41)	-0.00062* (-2.26)	-0.00072* (-2.79)	0.00011 (0.42)	0.0010* (1.97)
Party	–	-0.46* (-23.55)	-0.39* (-30.99)	-0.64* (-40.49)	-1.00* (-4.06)
Party* Days Since Introduction	–	0.00054* (3.14)	0.00070* (6.60)	0.0014* (12.44)	0.0050* (4.89)
Lame-duck	-0.39* (-5.42)	-0.22* (-5.40)	-0.15* (-5.22)	-0.31* (-8.76)	-0.17* (-2.96)
Lame-duck* Days Since Introduction	0.00042 (0.86)	0.00038 (1.17)	-0.00041 (-1.42)	0.00039 (1.47)	-0.000011 (-0.02)
Same Party as Sponsor	-0.23 (-1.44)	0.51* (29.30)	0.83* (6.76)	0.30* (7.84)	-0.12 (-0.48)
Same Party as Sponsor* Days Since Introduction	0.0085 (1.85)	-0.0027* (-18.79)	-0.0039* (-6.20)	0.00022 (0.87)	0.0052* (5.51)
On Considering Committee	0.89* (36.09)	0.73* (40.01)	0.59* (37.95)	0.66* (35.37)	0.41* (11.48)
On Considering Committee* Days Since Introduction	-0.0026* (-12.46)	-0.0020* (-11.64)	-0.0016* (-10.69)	-0.0015* (-9.40)	-0.0011* (-3.68)
Number of Cases	492,349	492,046	492,377	492,001	492,007
Chi-Squared	1306.27*	2905.35*	2597.93*	3623.53*	580.17*

* p < 0.05; Note: Cells marked “–” indicate coefficients not estimated due to collinearity.

Figure 1:
Probability of Cosponsoring Over Time Given Member-Proposal Match

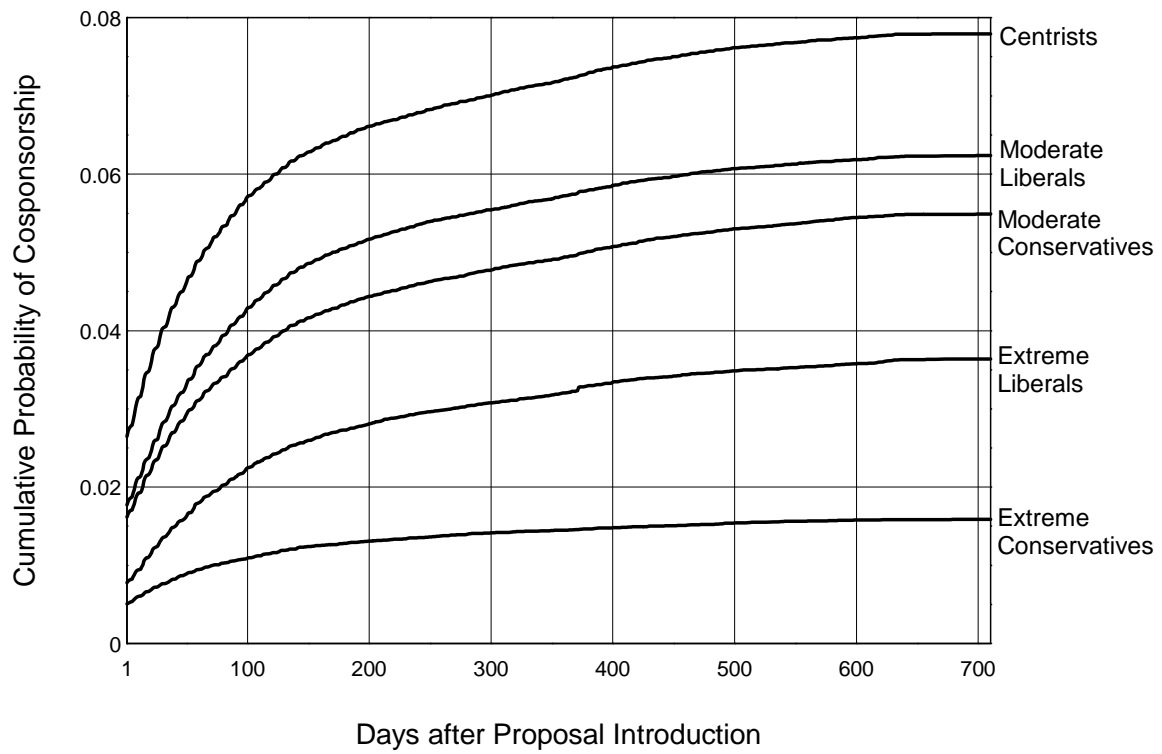


Figure 2:
Change in Cosponsorship Proportions Over Time
(Combining Liberals and Conservatives)

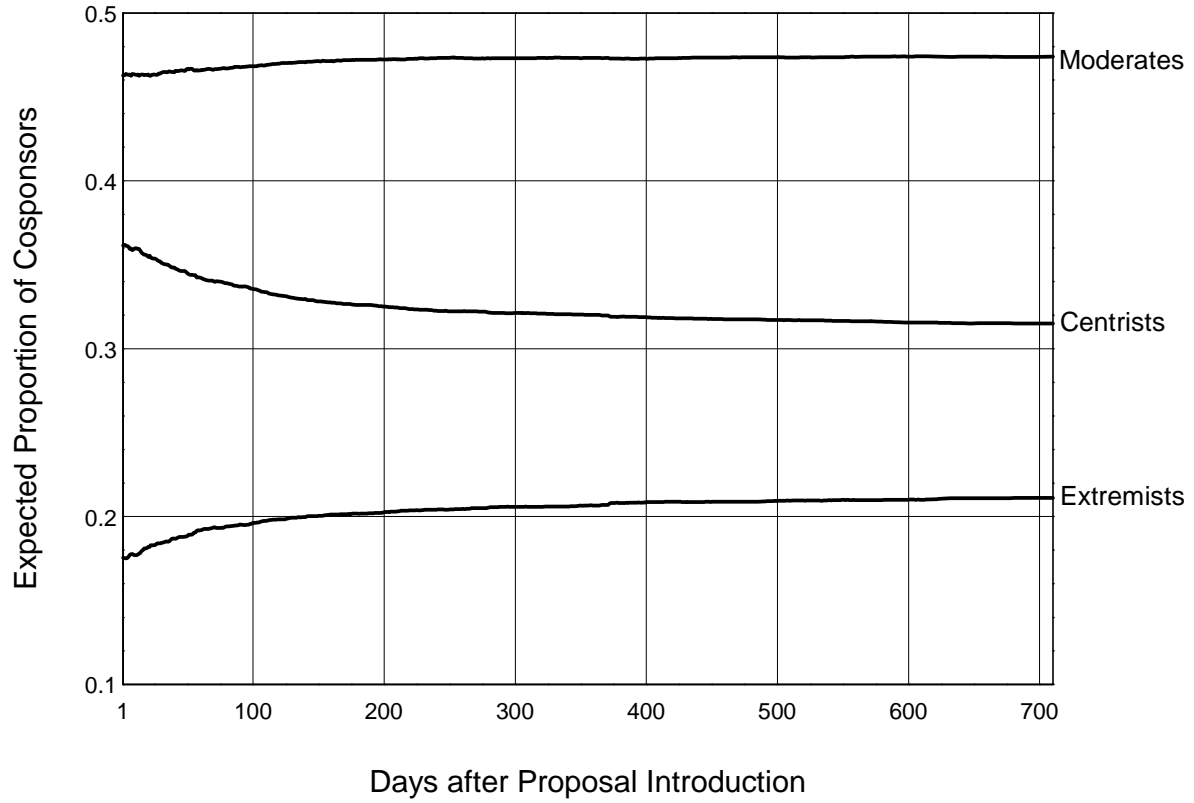


Figure 3:
Change in Cosponsorship Proportions Over Time for Proposals with 50 or More Cosponsors (Combining Liberals and Conservatives)

