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Uncovering the Hidden Effect of Party

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Some recent scholarship affords political parties little role in explaining patterns of legislative outcomes. Policy preferences, rather than partisanship, are said to provide the superior account of legislative behavior. In this paper, we challenge one recent such account of legislative outcomes. We show that the likelihood of finding a party effect depends on where we look for it and with what measures we use to test for it. Party effects, we find, are amply visible in the 1994 "A to Z" discharge petition campaign in the U.S. House of Representatives, a case where party has been termed inconsequential.

> Theories of legislative politics with a party component—while perhaps more realistic than their more parsimonious non-partisan counterparts—are not necessarily superior predictors of observable legislative behavior. (Krehbiel 1993, 237)

With these words, Keith Krehbiel poses a challenge to adherents of party-based theories of legislative politics. Here, as elsewhere (Krehbiel 1995, 1998), he suggests and marshals empirical support for a theory of legislative behavior in which explanatory power resides in legislators' policy preferences, not in their partisan affiliations. Contrary to assessments of Congress that grant key explanatory power to the homogeneity of preferences within legislative parties (Aldrich and Rohde 1995, 1997; Cooper and Brady 1981; Cox and McCubbins 1993; Rohde 1991; Sinclair 1995), Krehbiel suggests that partisanship has little to no independent effect on legislative behavior. Throwing down the gauntlet to legislative scholars, he sums up his challenge succinctly: "Where's the Party?" (1993, 235).

We argue here that it is premature to reject the hypothesis that majority party leaders can exert an independent effect on the behavior of their caucus members.

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We reach this conclusion by reexamining the 1994 "A to Z" discharge petition campaign in the U.S. House of Representatives, a case in which preferences rather than partisanship are said to provide the superior account of legislative behavior (Krehbiel 1995). We draw from the same spatial model used by Krehbiel, explore the conditions most likely to reveal significant party behavior, scrutinize the properties of two alternative measures of preferences, and show that party effects are indeed visible in the A to Z case. Legislative theories, we conclude, may in fact need to incorporate a partisan element.

Party versus Preferences in the Legislative Setting

We start with the nonpartisan theories cited by Krehbiel: those of Snyder (1991) and Groseclose (1996). Together, these models attempt to show how and why legislative leaders choose their targets in building coalitions. In short, legislators' preferences, not their party attachments, are said to be critical in identifying pivotal legislators in the eyes of legislative leaders. Assuming a simple unidimensional array of members' policy preferences, an optimizing leader is likely to expend resources on securing the votes of members approximately indifferent between the status quo and a proposed policy. Targeting fellow partisans, independent of their policy preferences, would be suboptimal in such a model. Greater side payments would be required to reach such "extremists" than would be required to secure the support of more proximate moderates, regardless of their party affiliation.

Figure 1 (adapted from Krehbiel 1995) makes clear the two options faced by legislative leaders in building a coalition: build a bipartisan coalition, based solely on members' policy concerns, or build a purely partisan coalition, overcoming distant partisans' policy preferences with partisan incentives, sanctions, and appeals. Numerous legislative scholars have argued that even relatively weak majority party leaders retain an arsenal of party-specific pressures that can be used to attract the support of fellow partisans (Aldrich and Rohde 1995, 1997; Cox and McCubbins 1993; Sinclair 1995). Majority party leaders might in fact find it cheaper to target fellow partisans instead of moderate members of the opposite party—as party leaders possess procedural and financial resources attractive to their party members. In short, a partisan alternative to the Snyder/Groseclose models suggests that party leaders will target fellow partisans over moderates of the opposing party. A test of the two theories condenses to a single question, as noted by Krehbiel (1995, 916): "Controlling for preferences, are Democratic leaders better able to attract support from their own party members than from Republicans?"[1]

[1] For a discussion of the importance of testing competing explanations, see Green and Shapiro 1994, which argues that rational choice theorists tend to "defend their favored sufficient explanations of known facts, without reference either to credible alternatives or to novel predictions" (183). Krehbiel (1995) avoids this problem by explicitly pitting two explanations against one another.

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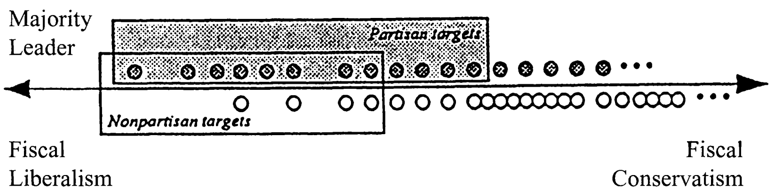
FIGURE 1

Alternative Partisan and Nonpartisan Coalitions of Legislative Leaders (Based on Krehbiel 1995)



Krehbiel (1993) himself has noted, however, that a joint test of hypotheses generated by nonpartisan and partisan legislative theories is not so simple. Consider the case where the majority party is highly cohesive and thus appears able to divert outcomes away from the preferences of the floor median and in the direction of the majority party median. Paradoxically, it is precisely this condition under which it is difficult, if not impossible, to detect an independent effect of party on members' behavior (particularly if we focus on members' roll call behavior). Party and preferences are mutually reinforcing influences in such cases.

Reproducing Krehbiel's (1993) depiction of members' ideal points under such conditions makes the case clearly (see Figure 2A). Here, the parties split cleanly on the policy choice. As Krehbiel argues, we could infer either that parties are strong because of high intraparty agreement or that members are voting consistently with their policy preferences, making parties irrelevant to the vote. "In spite of the cleanliness of the data in this example," Krehbiel points out, "the data cannot discriminate between a party hypothesis and a preference hypothesis" (1993, 238).[2]

When intraparty homogeneity declines, we have a much better chance to distinguish empirically between preferences and partisanship. Here, the alignment of preferences more closely resembles that depicted in Figure 2B, in which there is greater variance within each party. If the parties split cleanly on the vote in Figure 2B, some party members are clearly voting against their policy preferences. To the extent that we can detect an independent effect of party in such cases, we could conclude that party leaders have exerted influence over party members to elicit compliant party-line behavior when their policy preferences dictate otherwise. The policy outcome would be closer to the majority party median than the floor median in such cases. This is precisely the outcome that most

[2] The statistical problem that creates the inability to discriminate between the two hypotheses is collinearity between party and preferences. Symptoms of collinearity include: parameter estimate sensitivity to small changes in the data, individual coefficients for the collinear variables with large standard errors and large p values that are jointly significant, and coefficients with the "wrong" sign (Greene 1990, 279).

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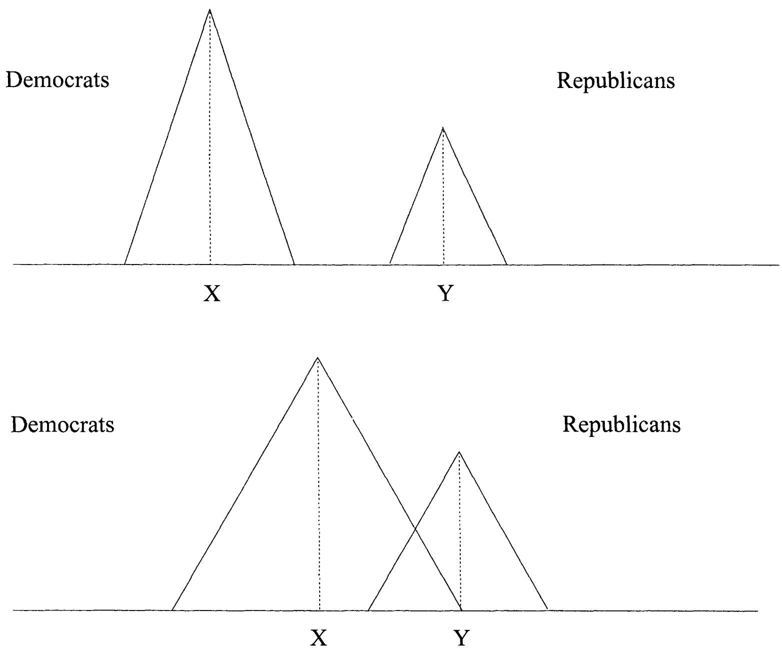
FIGURES 2A AND 2B

Two Hypothetical Distributions of Members' Ideal Points, by Party (Based on Krehbiel 1993)



party theorists predict. Rather than seeking out members (regardless of party) whose ideal points place them closest to the party median, majority party leaders would use their arsenal of procedural tools and their limited set of sanctions and rewards to encourage caucus members to support collective party goals (Cox and McCubbins 1993). Where successful, we would see an independent effect of party on legislative behavior, even after controlling for members' preferences.

Krehbiel's analytic demonstration is central to empirical analysis that attempts to test nonpartisan and partisan legislative theories. If party and preference are nearly perfectly aligned, it will be nearly impossible to test for the independent effects of party and preference. Indeed, as we demonstrate below, such collinearity is particularly likely to arise if an interest group has an incentive to disguise the partisan cast of its agenda in constructing its scores for members of Congress. In such cases, investigation of alternative measures that reflect the true distribution of preferences is warranted.

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The Measurement of Party and Preference

During the 103d Congress (1993–94), Representatives Robert Andrews (D-NJ) and Bill Zeliff (R-NH) cosponsored a deficit reduction bill that became known as "A to Z," a bill that would have given any member of the House the opportunity to offer floor amendments reducing outlays for any federal program, including entitlement programs. As such, the bill would have completely undermined the leadership's control over the floor agenda, one of its key sources of power (Bach and Smith 1988; Cox and McCubbins 1993; Sinclair 1995). Although a majority of the House cosponsored the bill, it was never reported from committee. When the bill's sponsors tried to circumvent the committee with a discharge petition, 33 members who had cosponsored the bill "waffled" by refusing to sign the discharge petition. Because only 203 House members had signed the petition by the end of the congressional session in October 1994 (15 members shy of the 218 required to discharge a bill), the wafflers killed the bill, the outcome preferred by the majority party leadership. According to press accounts, the discharge motion was defeated because the House Democratic leadership aggressively tried to discourage Democrats from signing the petition (Hager 1994; Kamen 1994; Pianin 1994a; Washington Post 1994; Will 1994).

Based on the Snyder (1991) and Groseclose (1996) models, Krehbiel hypothesizes that the decision to "waffle" on the A to Z bill "should be negatively associated with preference extremity and unaffected, at the margin, by majority party membership" (1995, 906). In equation (1) of Table 1, we replicate Krehbiel's analysis of the decision to waffle among cosponsors (1995, 920, col. 5). The dependent variable in this model (the decision to waffle) is derived from a June 17, 1994, *Wall Street Journal* editorial that listed the wafflers. Like Krehbiel's model, ours includes variables tapping members' policy preferences (***National Taxpayers Union*** and ***Americans for Democratic Action*** ratings), institutional interests (***Appropriations*** and ***Budget Committee Membership***), electoral security (***Electoral Margin*** and ***Seniority***), and party membership (coded 1 for ***Democrats,*** 0 otherwise).[3] As Krehbiel reported, policy preferences, money-committee membership and lower seniority—rather than party membership—appear to account for the waffling decision in June. Based upon this finding, Krehbiel concludes with a "tentative conjecture": "Perhaps legislative leadership has less to do with intra-party politics and more to do with inter-party or nonpartisan coalition-building than recent studies suggest" (1995, 922).

There is an important caveat, however, to such an interpretation of the data. Figure 3 shows the distribution of members' policy preferences based on ***NTU***

[3] Like Krehbiel, we code seniority as the year in which the member was first elected. Thus, the more senior the member, the lower the score.

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TABLE 1

Comparison of Party Effects on Waffling Behavior with Different Preference Measures, June 17, 1994 Discharge List

	(1)	(2)	(3)
Constant	6.976	5.390	3.794
	(3.028)	(2.714)	(2.239)
ADA	.00205	.0116	
	(.260)	(1.516)	
NTU	−.0513		
	(−2.517)		
Concord Coalition		−.0371	
		(−2.543)	
Deficit factor score			.886
			(2.782)
Democrat	.317	1.432	1.082
	(.438)	(3.294)	(2.417)
First elected (Seniority)	−.0673	−.0689	−.0606
	(−3.361)	(−3.158)	(−3.129)
Electoral margin	−.00129	−.00325	−.00326
	(−.257)	(−.600)	(−.616)
Appropriations member	1.111	1.236	.986
	(2.889)	(3.375)	(2.560)
Budget member	.867	.927	.726
	(1.549)	(2.017)	(1.301)
Number of Cases	226	226	226
Log Likelihood	−54.32	−57.31	−56.56
Percent Correct	88.94	87.61	88.94
Reduction in Error	21.98	12.50	21.88

Probit estimates with asymptotic t-statistics shown under coefficients. "Robust" (Huber/White) standard errors were used for the calculation of t-statistics. Equation (1) essentially replicates Krehbiel's results (1995, table 4, col. 5). The insignificant discrepancies stem from three sources: our use of robust standard errors, our correction of a few members' ADA scores, and our dropping of two members who were elected in the middle of 1993: Portman (R-OH) and Barca (D-WI). Although the NTU calculated scores for Barca and Portman, they were not in office for roughly half the votes used in the NTU index. We dropped them for comparability with our results presented in equations (2) and (3) because the Concord Coalition did not calculate scores for either member.

scores, separated by party.[4] The pronounced bimodal distribution of the ***NTU*** scores resembles the depiction of party and policy preferences in Figure 2A.[5] As suggested by Krehbiel two years earlier, such an alignment of preferences "makes it impossible to discriminate between a simple and parsimonious preference-based theory and a more complex and elaborate preference-and-party

[4] Figures 3 and 4 are kernel density plots generated by the "kdensity" procedure in Stata 6.0. A kernel density plot is essentially a smoothed histogram. Rather than plotting a series of bars to show the number of members with either NTU (Figure 3) or Concord Coalition (Figure 4) scores, these

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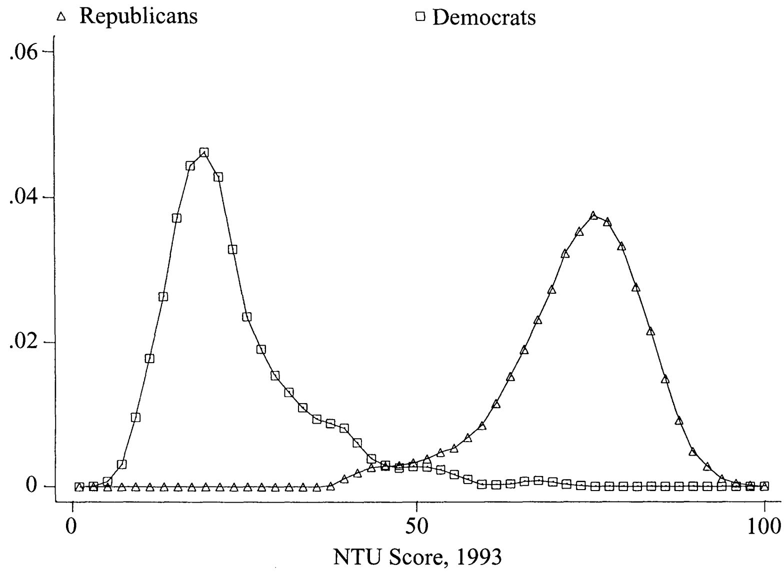
FIGURE 3

Kernel Density Plots of 1993 NTU Scores, by Party



See note 4 for a detailed explanation of kernel density plots.

theory" (1993, 262). Indeed, the correlation between NTU and party is extremely high (−.93). Among cosponsors, the correlation is −.85. Such high correlations make it near impossible to sort out the independent effect of party and policy preferences with such measures.

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figures report empirical probability distributions of the interest group ratings for both parties. To smooth the plot, the procedure calculates the likelihood that a member will fall at each point on the scale by averaging the likelihood that a member will fall in an interval surrounding each point. There are several ways of calculating this average. The plots contained in figures 3 and 4 are based on the Epanechnikov kernel function. Alternative kernel functions yield similar results. The advantage of a kernel plot over a histogram is that it is easier to read than the 435 separate bars that would be needed to present Democratic and Republican histograms. See Fox 1990 for a discussion of kernel density estimates. The plots should be interpreted the same as a histogram. For NTU, the Democratic mean is 23.38 and the standard deviation is 10.25; for Republicans, the mean is 72.81 and the standard deviation is 9.63. For Concord, the Democratic mean is 39.56 and the standard deviation is 8.70; for Republicans, the mean is 52.65 and the standard deviation is 7.81. Because the Democratic caucus contains a few members with extreme scores, the Democratic scores span a broader range than Republican scores.

[5] For a more complete discussion of extremism and bimodality in interest group ratings and the possible statistical consequences, see Snyder 1992.

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If NTU scores were the best reflection of the true distribution of members' preferences on deficit reduction, we would not be able to statistically disentangle the effects of party and preferences. Scrutiny of the NTU, however, reveals that their scores poorly capture the distribution of such deficit reduction preferences. Most importantly, the high correlations between party and NTU scores appear to stem from the NTU's distinctly partisan agenda. Contrary to NTU's claims to be a nonpartisan lobbying organization (NTU n.d.), evidence suggests that the NTU strongly prefers Republicans and that its preference for Republicans affects its methodology of calculating NTU scores for members of Congress.

First, consider the response of NTU Executive Vice President David Keating to election outcomes in the midterm elections of 1996: "I'm disappointed that in these races between people who were for us and who were against us, the people who were against us won. Every close race in the Senate has gone to the Democrats" (as cited in Grove 1996). Second, the political action committee of the avowedly nonpartisan NTU reported to the Federal Election Commission that it gave $2,000 in February 1994 to the Republican National Committee's soft money account.[6]

Third, and most important for our study, the methodology for creating members' NTU scores seems crafted in part to reward Republicans and punish Democrats. To be sure, others have documented biases in interest group ratings of members of Congress (see, e.g., Fowler 1982; Smith 1995). In this case, the group's friends and enemies appear decidedly partisan. For example, the NTU sometimes excludes votes that would make Democrats look better in their rankings.[7] In addition, after selecting the votes to be included in members' scores, the NTU assigns weights (from 1 to 100) to the votes according to the "relative effect of each vote" on total federal spending and the "political effect of a vote" on future spending (NTU n.d.).[8] Analysis of the 1993 vote weights suggests that on the most heavily weighted votes Republicans voted overwhelmingly for the

[6] Apparently beholden to Republican Party interests, the NTU also withdrew its support in 1995 for cutting mining and grazing fee subsidies when its attacks "drew fury from House Republicans" (Shafer 1995; see also Kamen 1995). The NTU reportedly withdrew its support for reform when its prime contributor threatened to withdraw his funding (see Kuntz 1995). That contributor, Richard Mellon Scaife, has been called "the conservative movement's most valuable asset" and is a well-known backer of Republican causes (Kuntz 1995).

[7] On November 22, 1993, the House voted on a series of amendments to a "reinventing government" measure. All of the amendments would have reduced spending by varying amounts. The Sabo amendment, backed by 96.1% of Democrats and the Clinton administration but opposed by 79.0% of Republicans, would have cut spending by $37.1 billion over five years. Strangely enough, the Sabo amendment was not included in the NTU ratings, even though two other amendments (Penny-Kasich and Frank-Shays) from the same day on the same measure were included. The partisan skew of NTU scores are discussed further in Shear 1994.

[8] In 1993, the NTU ratings were based on 271 votes, but most were not heavily weighted; more than 100 received a rating of either one or two.

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pro-NTU position and Democrats voted overwhelmingly against the NTU.[9] Even a former chair of the NTU has assailed the partisan skew of its ratings: "It seems that NTU's ventures into partisan politics reflect incredibly bad judgment of where the average taxpayer's interests lie and require grotesquely illogical rating measures to rationalize the misjudgments" (Fitzgerald n.d.).

To the extent that party leadership or party loyalty affects votes that are weighted heavily in the NTU ratings, when NTU and party are both used as predictors (as in Table 1, equation 1), NTU will explain some of the variance more properly accorded to party. In sum, by omitting votes that favor Democrats and weighting partisan or near party-line votes more heavily, the NTU ratings exaggerate the differences between the two parties (contributing to the bimodality seen in Figure 3), obscure the true distribution of members' preferences on deficit reduction, and make it more difficult to measure the separate effects of party and preferences.

These findings lead us to substitute an alternative measure of members' preferences on deficit reduction: scores created by the nonpartisan Concord Coalition. Unlike the NTU, which was chaired in the 103d Congress by its founder, James Dale Davidson (a conservative activist with ties to the Republican Party), the Concord Coalition was run at that time by its founders, former senators Paul Tsongas (D-MA) and Warren Rudman (R-NH). As shown in Figure 4, the ***Concord Coalition*** scores produce far less bimodally distributed preferences between the two parties, correlating with party at $-.61$ for all members and at $-.31$ for cosponsors. Whereas the distribution of ***NTU*** scores more closely resembles the hypothetical distribution in Figure 2A, the distribution of ***Concord Coalition*** scores resembles Figure 2B—a distribution that enables us to test more easily for significant party behavior. Substituting ***Concord Coalition*** scores for ***NTU*** scores, we rerun the analysis, with results shown in Table 1, equation (2). Now, both party and policy preference coefficients are statistically significant. It appears that intraparty politics has an important role in the decision to waffle.

Our results are *not* a product of substituting an interest group score favorable to our case for one hostile to our case. That is, our substantive conclusions are not merely the result of a subjective choice of Concord scores over NTU scores. To check the face validity of the Concord Coalition scores and the robustness of our results, we also create a ***Deficit Reduction Factor Score*** by factor-analyzing four scores: ***NTU, Concord Coalition, ADA,*** and ***NOMINATE*** scores (the latter two being general ideological measures). Clearly, a measure constructed from these four scales is a more valid and reliable measure of members' preferences than

[9] For example, in the 1993 ratings nine votes received weights of higher than 35. These votes were mainly on budget reconciliation and other big budget bills. On these nine votes, the average vote had 91.3% of Democrats taking the anti-NTU position and 92.2% of Republicans taking the pro-NTU position. On four of these votes, not a single Republican member voted against the NTU-favored position.

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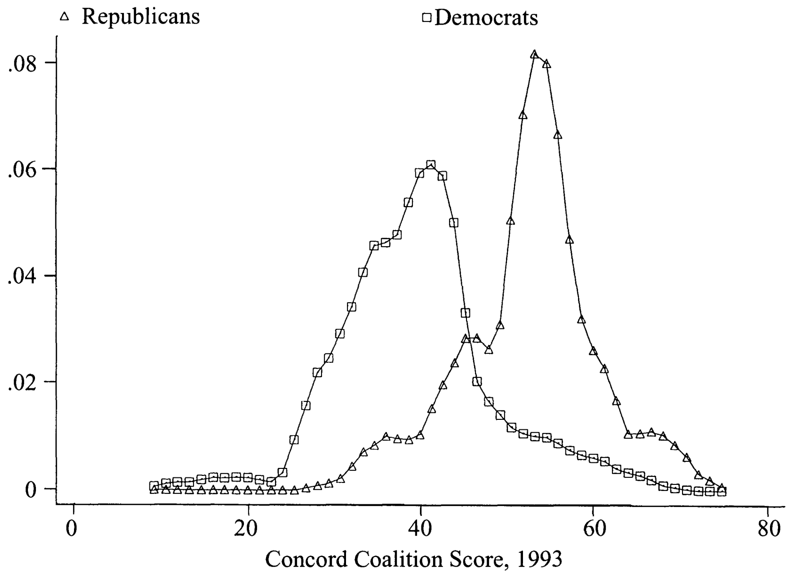
FIGURE 4

Kernel Density Plots of 1993 Concord Coalition Scores, by Party



See note 4 for a detailed explanation of kernel density plots.

the partisan-tainted NTU scores.[10] By the same logic, the ***Factor Scores*** also reduce the effect of idiosyncrasies that may be present in the Concord scores. Substituting the ***Factor Scores*** in place of the raw Concord Coalition scores as measures of members' preferences (Table 1, equation 3) yields the same substantive results: both party and preferences are statistically significant in predicting waffling decisions in June. Because we obtain similar results with both ***Factor Scores*** and ***Concord Coalition*** scores alone, we are confident that both party pressure and policy preferences influence members' waffling decisions.[11]

[10] The factor scores were created by a principal factors analysis with a varimax rotation. NTU, Concord, Nominate (first dimension), and ADA scores have zero-order correlations with the deficit reduction factor of −.95, −.86, −.81 and .69 respectively. The higher correlations of the NTU and Concord scores provide face validity to our interpretation that the factor is indeed tapping fiscal ideology, rather than simply a more general left–right orientation. Using the shared variance of four correlated indicators of underlying deficit reduction preferences, rather than simply one indicator alone, increases the construct validity and the reliability of our measure of deficit reduction preferences, giving us confidence that we are more accurately assessing the relationship between deficit reduction preferences and waffling behavior (Carmines and Zeller 1979; Nunnally and Bernstein 1994, chaps. 2 and 3).

[11] Substituting Concord scores in place of NTU scores also affects the anomalous results of party effects on cosponsorship (Krehbiel 1995, 911). In Krehbiel's model including NTU, ADA, and party,

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Uncovering Further Effects of Party

The use of NTU scores is not the only reason that the effects of party might have been obscured in Krehbiel's (1995) analysis. Recall Krehbiel used a list of wafflers from a June 17 *Wall Street Journal* editorial. The *Journal,* however, published complete lists of wafflers in editorials on May 24, June 17, and October 12, 1994. Although the June 17 coalition was essentially stable, the May 24 coalition was not. Of the 58 who waffled when the May editorial appeared, 28 (48.3%) signed the discharge petition by June 17. A close scrutiny of the events between May 24 and June 17 suggests analysis of the May list of wafflers might produce a different set of results.

Between May 24 and June 17, two sets of negotiations occurred in the House. One group, led by conservative Charlie Stenholm (D-TX), was negotiating with the A to Z sponsors to secure changes in the bill. A deal was announced on June 14, and 19 Democrats (including Stenholm) signed that day.[12] At the same time, a second group of A to Z cosponsors, led by Bill Orton (D-UT), was negotiating with the Democratic leadership. This group promised not to sign the discharge petition if the Democratic leadership would make a commitment to consider an alternative entitlement reform bill on the House floor. After the Stenholm deal was announced, *Congress Daily* quoted Orton as saying that a deal with the leadership was imminent. On June 16, *Congress Daily* reported that the leadership was "optimistic" that A to Z supporters would have "trouble obtaining the final signatures." By the time the *Wall Street Journal* came out the next morning, six more members had signed the discharge petition, bringing the total number of signatures to 203. There the discharge drive stalled, the leadership having agreed to bring up entitlement reform as the price for securing enough Democratic commitments to defeat the petition. The next day, Chief Deputy Whip Bill Richardson (D-NM) was quoted as saying: "We feel good that we've stemmed the erosion" (Hager 1994).

In modeling cosponsors' discharge decisions, the dependent variable used by Krehbiel is the list of wafflers that appeared in June after the Stenholm and Orton deals had been cemented. But what if he had used the May 24 list? Judging from

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the coefficient on party is statistically significant, but suggests that Democrats were *more* likely to cosponsor A to Z once one controls for preferences. This is a startling finding, given that 97% of Republicans but only 22% of Democrats cosponsored the measure. This result appears to be an artifact of the extremely high correlation of −.93 between party and NTU (among all members). When the Concord scores are substituted for the NTU scores in Krehbiel's original cosponsorship analysis (1995, Table 2, col. 5), the marginal effect on cosponsorship of being a Democrat goes from 31.3% *more* likely to 39.8% *less* likely. Therefore, when using the Concord Coalition scores with their lesser collinearity problems, one can conclude that party effects were present at both the cosponsorship and waffling stages. The corresponding marginal effect of being a Democrat when factor scores are used is 40.8% less likely to cosponsor.

[12] *Congress Daily* reported June 13 that 178 members had signed. On June 15, *Congress Daily* reported that 197 had signed.

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the chief deputy whip's June comment, if party effects are visible, they would most likely appear *before* the Orton and Stenholm deals were concluded. As noted above, Democratic leaders had tried to discourage Democrats from signing the petition; but once Democratic leaders had secured the support of the Orton group, they were confident that they had beaten the discharge petition. They might then have eased up their pressure on the remaining waffling Democrats, thereby attenuating the relationship between waffling and party. Alternatively, party leaders might have days earlier lessened their pressure on the Stenholm coalition not to sign since they believed a deal with Orton was imminent.[13] Such behavior by party leaders under either scenario would be consistent with the empirical regularity of "pocket voting" observed by numerous students of Congress (see, e.g., Aldrich and Rohde 1997, 15; Froman and Ripley 1965, 55–56; King and Zeckhauser 1997; Sinclair 1995, 247) and occasionally by members themselves (O'Neill and Novak 1987, 134). Party leaders secure the commitments of their caucus members to withhold taking a position (here, pressuring them to waffle), until it can be determined whether their support is needed. Using the June list, in other words, might camouflage partisan effects visible before negotiations had concluded.

The phenomenon of pocket voting has often been noted, but has never been rigorously tested in a multivariate fashion. Our intuition from the logic of pocket voting suggests that the coefficient for party should be statistically significant in May before the Orton and Stenholm deals were reached. Thus, in Table 2, we use the only available list of wafflers (*Wall Street Journal*, 24 May) as our dependent variable. In all three of the May models, party effects are clearly visible and in the predicted direction. Even if ***NTU*** scores are used to control for policy preferences (Table 2, equation 1), ***Democrats*** are still more likely to waffle in May (with a predicted probability change of .496, a very large effect).[14] Although more conservative members (based on their ***NTU*** scores) are at this earlier stage less likely to waffle (−.077)[15] and ***Appropriations Committee*** members and

[13] Andrews had suggested as much in *Congress Daily* on June 15.

[14] The predicted probability changes were calculated as follows. For the party effects, the predicted probabilities were calculated for Democrats and Republicans, holding all other variables at the cosponsor sample means. For the interest group ratings, we calculated the marginal probability changes of a one-standard-deviation change in the interest group rating, again holding all other variables at the cosponsor sample means. Our calculations of the predicted probability changes of the interest group scores are much lower than those reported by Krehbiel (1995, 919). For example, Krehbiel calculates the effect of a one-standard-deviation shock to NTU ratings as changing the probability of waffling at −.370, while we calculate it to be −.033. This difference occurs because our baseline probabilities are calculated only for those members who cosponsored the bill—i.e., those members in the sample used to calculate the parameter estimates. Krehbiel instead uses the full House sample to calculate his baseline probabilities. The magnitude of the NTU effect we calculate is still relatively "large," since only 14% of the sample waffled.

[15] It is important to note that although the magnitude of the NTU effect in May (−.077) is larger than that in June (−.033), there is more "room to move" for a negative effect in May, as 25.2% waffled in May but only 14.2% waffled in June.

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TABLE 2

Comparison of Party Effects on Waffling Behavior with Different Preference Measures, May 24, 1994, Discharge List

	(1)	(2)	(3)
Constant	4.154	2.724	2.583
	(2.423)	(1.434)	(1.570)
ADA	.00419	.00961	
	(.547)	(1.373)	
NTU	−.0254		
	(−1.706)		
Concord Coalition		−.00810	
		(−.556)	
Deficit factor score			.328
			(1.225)
Democrat	1.696	2.287	2.358
	(3.265)	(5.882)	(5.628)
First elected (Seniority)	−.0512	−.0515	−.0489
	(−2.667)	(−2.531)	(−2.560)
Electoral margin	.00355	.00309	.00237
	(.589)	(.504)	(.386)
Appropriations member	.726	.877	.720
	(1.686)	(2.161)	(1.737)
Budget member	.557	.496	.491
	(1.174)	(1.188)	(1.017)
Number of Cases	226	226	226
Log Likelihood	−61.86	−63.66	−64.00
Percent Correct	89.38	89.38	89.38
Reduction in Error	57.89	57.89	57.89

Probit estimates with asymptotic t-statistics shown under coefficients. "Robust" (Huber/White) standard errors were used for the calculation of t-statistics.

more senior members are still more likely to waffle, the level of significance on the NTU coefficient is slightly diminished from June results.

Substituting the ***Concord Coalition*** scores and the ***Deficit Reduction Factor*** scores for the ***NTU*** scores (Table 2, equations 2 and 3), we find strong support for the independent effect of party on members' behavior and little support for the non partisan vote-buying/favor-trading theory: Democrats are more likely to waffle early in the discharge campaign (with a change in probability for being a Democrat of .685 and .707 for equations 2 and 3, respectively). Using either alternative to the ***NTU*** scores, the coefficient for policy preferences is no longer statistically significant.

Although party is significant in both the May and June models when ***Concord Coalition*** scores or factor analytic scores are used (Tables 1 and 2, equations 2 and 3), the effect of party membership is substantially attenuated in the June model—precisely what we would expect if Democratic leaders had by mid-June

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already secured enough pocket commitments from Democrats to waffle, thereby ensuring defeat of the discharge campaign. Knowing that the bill would not be dislodged from committee, other Democrats were then free to sign the discharge petition—allowing factors such as policy preferences to shape their salient procedural choice. In other words, the data are consistent with the idea that the Democratic leadership in May had been able to pressure its members to withhold their signatures from the discharge petition.

The fate of the A to Z bill suggests that Democratic leaders were successful in diverting the policy outcome away from the preferences of the median voter (as expressed by the cosponsorship of the A to Z bill). A criticism might be raised, however, that any such party effects were still inconsequential, as majority party leaders were ultimately required to make policy concessions to secure the support of the Orton coalition. This argument is only credible if the alternative entitlement bill offered to Orton was a genuine policy concession. We find no evidence that this was the case. The procedural motion to bring the entitlement bill to the floor was defeated, 83-339, by a bipartisan coalition. During the debate over the rule, Newt Gingrich (R-GA) explained:

> Let us be clear that this rule and this resolution is a case study in how the liberal Democratic leadership manipulates the House. . . . Under the new open discharge petition rule . . . we were actually on the verge of getting enough votes that we could actually bring the A-to-Z spending cuts to the floor. . . . At this point, the liberal Democratic leadership went in overdrive, and they began to give you a smoke screen to go home and claim you accomplished something. . . . This bill . . . is a classic example of how the liberal Democratic leadership picks off one element of their party at a time to maintain control of the House against the will of the American people (Gingrich, *Congressional Record,* 5 October 1994, H-10854).

Other less partisan observers reached similar conclusions (Pianin 1994b). In the end, the legislative outcome on A to Z was consistent with the goals of the majority party: the status quo favored by the median member of the majority party prevailed.

Conclusion

> [T]he concepts [issues and factions] are not identical and must be distinguished. They have been unnecessarily combined and confused in legislative analysis, especially roll-call analysis. The same data, used in similarly computed and highly correlated indices, have been used by different authors to infer legislators' "liberalism–conservatism" and "partisanship." The condition under which one or the other use of such data is permissible have not been adequately explored. (MacRae 1970, 6–7).

As suggested by MacRae nearly 30 years ago, unraveling the effects of party and preferences is not simple. As Krehbiel has pointedly shown, it is not enough to show that a strong majority party achieved the policy outcomes preferred by its members. "Politics," Krehbiel suggests, "should be significantly different with parties from what it is without them" (1993, 240). Clearly in the A to Z case, politics without parties should have led to the discharge of A to Z. A majority of

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the House had cosponsored the bill, so a majority of the House should have signed the discharge petition. That did not happen. Both journalistic accounts and statistical analysis suggest that party leaders targeted fellow partisans in seeking to derail the discharge campaign. The A to Z outcome cannot be accurately explained with recourse to a nonpartisan theory.

Although we are reluctant to draw generalizations from a single case, the differences that result from substituting ***Concord Coalition*** scores and ***Deficit Reduction Factor*** scores for ***NTU*** scores reinforce the difficulty of identifying accurate measures of members' policy preferences (see Bailey and Brady 1998). Indeed, as Epstein and Mershon have argued, "when analysts adopt surrogates of actors' political preferences for purposes unanticipated by the inventors of those measures, they often stretch . . . the range of reliability and validity" (1996, 261). The A to Z case also gives us solid ground for pursuing in greater detail a party theory of pocket voting or pocket commitments. Our ability to observe the effects of party will likely require a more complex and dynamic partisan theory. Richard Fenno suggested as much when he noted:

> If we are to explain outcomes, who decides when may be as important to know as who decides what. We have devoted more energy to studying policy positioning in space than to studying policy sequencing in time. To our rich comprehension of the politics of left, right and center, we can usefully add an equally rich comprehension of the politics of early, later, and late. (1986, 9)

Our efforts here to build on Krehbiel's (1993, 1995) analytic and empirical results suggest both the difficulty and importance of deriving and testing a more detailed and analytically robust theory of party effects.

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