District Partisan Safety and Constituent Evaluations of U.S. House Members

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Disputing the conventional wisdom of congressional scholars, Thomas Brunell has proposed that drafters of congressional district lines should strive to create the maximum number of safe partisan seats. One major reason, to which he devotes considerable attention, is that more votes will be cast for incumbent winners in more homogeneous districts and, because voting for a winning candidate arguably elevates the esteem in which the incumbent is held, district opinion of safer incumbents should therefore be higher. In my own study, I find that district homogeneity, in fact, only modestly improves incumbent positivity. Part of the explanation seems to be that opposition party identifiers, while less abundant in safer districts, have disproportionately critical views of safer members, likely because of these members' ideological extremity. Moreover, I uncover only mixed evidence supporting Brunell's assertion that the act of voting for a victorious incumbent has an independent effect in raising post-election popularity.

Certainly, one of the most important preoccupations of congressional election scholars in recent years has been the emergence of increasingly polarized U.S. House districts. With growing frequency, the median district voter now occupies a position considerably off to the blue or red side of the partisan divide. Theriault, for example, shows this forcefully by comparing the number of districts in 1976 and 2004 that delivered at least 60 percent of their vote to either party's presidential candidate. Despite the near-identity of the winner's national popular vote margin in the two elections—approximately 2.0 percent for Carter and 2.5 percent for Bush—the number of districts meeting the 60 percent threshold almost doubled from 113 to 217 across the 28 year period (2008, 3-4).[1]

Attempts to identify the causes of recent district polarization have triggered a lively debate, centered around the respective roles played by redistricting (Mann 2005, 93; Oppenheimer 2005, 149-52; Abramowitz et al. 2006, 78-79; McDonald 2006, 227-33), growing reliance upon cultural and social affinity with prospective new living areas when Americans make relocation decisions (Oppenheimer 2005, 152-53), and voters bringing their partisanship into alignment with their ideology through the "sorting" process (Levendusky 2004, 13-26; Abramowitz et al. 2006, 79-80).[2] But there is far less disagreement concerning the consequences of greater partisan homogeneity for representative government. Probably the most ubiquitous criticism is that non-competitive districts weaken members' responsiveness

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to the average general election voter, making them respond instead to relatively extreme voters who dominate primary elections (McDonald and Samples 2006, 6-7). Members not fearing election defeat likewise may be tempted to abuse their power for self-serving ends such as monetary gain (McDonald and Samples 2006, 4-5).

It is therefore somewhat of a surprise to come across the contrarian views about district safety and responsiveness recently expressed by Thomas Brunell, part of his extended argument in favor of creating the maximum number of secure electoral strongholds for each party. Brunell sees primary election competition taking the place of general election competition as the mechanism for making members stay responsive to constituents and refrain from ethical abuses. Primaries in more homogeneous districts tend to attract more competition and to have higher turnout (2006, 81; 2008, 12, 100).[3]

The most important representational reason for creating greater numbers of homogeneous districts, according to Brunell, is that aggregate public opinion in these districts should be marked by more positive attitudes toward the local Congress member and the institution of Congress, as well as higher levels of political efficacy. One full chapter of his book is devoted to the topic (Chapter 3), and it is here where most of his own statistical analysis is concentrated. In a period of widespread cynicism toward government, any political arrangement capable of producing the attitudinal improvement suggested by Brunell deserves to be given serious consideration. From a logical standpoint, voters supporting the victorious candidate should not only develop more positive political outlooks because of the vicarious pleasure inherent in winning per se, but because the successful candidate is more likely to be responsive to their interests (2008, 29).

From an empirical standpoint, Brunell is able to show that indeed, individual voters casting ballots for winning incumbents have higher scores on measures of incumbent and Congress satisfaction, and on political efficacy. By extrapolation, therefore, he concludes that districts with greater partisan homogeneity, simply because they include more voters who cast ballots for the winner, should be characterized by higher overall levels of satisfaction and efficacy. But this conclusion is purely inferential. Brunell himself never directly demonstrates that more homogeneous districts, in fact, have these characteristics.

At a more fundamental level, however, there is the question of how strong the effects of voting for winning incumbents actually are. Brunell conducts pooled cross-sectional analyses covering all elections from the ANES (American National Election Studies) Cumulative File 1948-2004 that contain the requisite data. Two models each are estimated when approval of Congress and efficacy are employed, respectively, as the dependent variable, but only one model per variable finds voters for winners to be at a significantly higher level than voters for losers (2008, 44). Stronger relationships emerge when incumbent House member evaluations are the dependent variable. But here, simply finding a strong relationship between pro-incumbent voting and liking the incumbent begs the question of causality. Does the act of voting cause incumbent approval? Or is it incumbent approval that causes the vote? Brunell's procedure does not permit an answer to this essential question.

Both questions I have just spotlighted—whether politically homogeneous and non-homogeneous districts differ in the hypothesized direction, and whether voting for winning incumbents generates more positive political attitudes—will be investigated in my own study. I have chosen specifically to focus on incumbent evaluations rather than on Congress appraisals or efficacy, because it is here where voting for winning incumbents has the strongest effect in Brunell's analysis and thus where his prescription for carving out the maximum number of homogeneous districts should produce the greatest change. The evaluative measures I employ are derived from all three types of questions regularly asked by the ANES to gauge summary appraisals of the incumbent: whether the respondent approves of the incumbent's job performance, feeling thermometer ratings of the incumbent, and the net positivity of the member in terms of open-ended comments volunteered about him or her. Analysis extends across the same period covered by Brunell; i.e., from 1978, the first year in which the ANES asked these questions, through 2004. (However, no question about job approval was asked in 2002; furthermore, the open-ended comment analysis must halt in 2000, because the relevant data do not exist in either 2002 or 2004.)

Differences between Homogeneous and Non-Homogeneous Districts

To provide a first look at the effects of partisan homogeneity, I dichotomize districts into those with greater or lesser degrees of safety and compare them with regard to constituents' evaluations of their members. Later, when I move beyond this preliminary examination to perform multivariate analysis, homogeneity will be considered in continuous terms. The measure of homogeneity, in accordance with standard practice, relies upon two-party presidential vote returns. In districts with a Democratic member, homogeneity is the district two-party vote proportion for the Democratic presidential candidate minus the mean proportion of the Democratic presidential vote occurring across all 435 districts. In districts with a Republican member, homogeneity is the mean proportion of the Democratic presidential vote across all districts minus the Democratic proportion in the district. Thus, in all cases more positive values of the measure indicate greater homogeneity. When a presidential election year is analyzed, I rely upon election returns from that same year. For midterm election years, the returns utilized are from the presidential election of two years earlier.[4]

To dichotomize districts in each election according to their amount of homogeneity, I use as the cutting point the mean value of partisan homogeneity that year calculated across all ANES respondents with members running for reelection.[5] The safer group of districts operationalized in this way has a level of homogeneity that is, on average, a sizable 16.2 percentage points greater than is the case in less safe districts. Coding of the incumbent evaluations to be compared across the two kinds of districts is as follows:

Incumbent job approval (1 if the respondent strongly or not strongly approves of the incumbent, 0 if strongly or not strongly disapproves)[6]

Thermometer rating of incumbent (rating of incumbent on 0°–100° feeling thermometer)

Net positivity of comments about incumbent (number of positive open-ended comments volunteered about incumbent minus number of negative comments).[7]

A final inter-district comparison—this time separate from incumbent evaluations—is performed in terms of the party identification of the respondents. This serves to provide a sense of how well the district presidential vote actually proxies the distribution of partisanship. Here, the coding scheme is:

Partisanship (2 if respondent is strong or weak identifier with party of incumbent, or independent leaner; 1 if pure independent; 0 if strong or weak identifier with party opposite that of incumbent, or independent leaner).

Table 1 presents the evaluation differences between districts and the corresponding t-tests when first, all respondents with non-missing data are analyzed, and then, just respondents casting House votes. The election-by-election means calculated for all respondents show weak to non-existent support for the hypothesis. Constituents from safer districts do approve of incumbent job performance more often than do those from less safe districts in nine of the 13 yearly comparisons, but only four of these differences are significant ($p<.05$, one-tail t-test). Thermometer rating differences are in the hypothesized direction just five of 14 times, with two being significant, while the comparable figures for the net positivity measure are three of 12 times with significance never reached. Restricting the data set just to House voters only marginally yields results more compatible with the hypothesis. For job approval, differences between safer and less safe districts are no more likely to be in the proper direction or to achieve significance than before. Seven differences in thermometer ratings and eight differences in net positivity now have the correct sign, with four and one, respectively, being significant.

In addition, I perform in Table 1 pooled cross-sectional analyses of incumbent evaluation differences by combining all cases falling within each competitiveness category together across elections. The consequence, once again, is not encouraging to Brunell's hypothesis. Safer districts are significantly more pro-incumbent only once in the six total comparisons involving all respondents and voters only, when job approval by voters is considered. With regard to the more elementary question of whether, at least, the direction of difference corresponds to expectations, only three comparisons show more positive evaluations existing in safer districts (i.e., job approval assessments by all respondents and by voters, and thermometer ratings by voters). A fourth effectively results in a tie (i.e., the net positivity of voters' comments).

At the same time, however, the two kinds of districts are very strongly differentiated on the basis of their constituents' party identification. Whether the election-by-election or pooled analyses are considered, every mean of the partisanship variable (which takes on larger values as more constituents affiliate with the incumbent's party) is greater in safer districts. And the differences likewise are significant in all but three instances (1986 with just voters analyzed, and 1990 with all respondents as well as just voters analyzed). So I now face the task of resolving the apparent paradox of why, if respondents in safer districts are much likelier to identify with the party of their incumbent representative, they have only a modest tendency to more positively evaluate that incumbent.

One possibility is that while there are fewer independents and opposition party identifiers in safer districts, their views of the incumbent may be considerably negative, especially those of opposition party identifiers. This could result from the fact that incumbents from more homogeneous districts are more ideologically extreme (Mann 2006, 275-79; Erikson and Wright 2009, 83-85).[8] Electorally buttressed by the partisan makeup of their constituencies, they have little incentive to throw roll call votes in the direction of non-incumbent party respondents (Oppenheimer 2005, 154-55). Thus, these respondents may find them quite unattractive, and such pronounced negativity might offset in part the evaluative boost afforded incumbents in safer districts by the sheer number of incumbent party identifiers residing there.

I test out this possibility via a three-step process that focuses on the role played specifically by opposition party identifiers. First, with regard to each of the three kinds of member evaluations, I see whether opposition party

Table 1. Mean Incumbent Evaluations in Safer and Less Safe House Districts

	Job Approval		Thermometer Ratings		Net Positivity		Partisanship	
Year	Safer District	Less Safe District	Safer District	Less Safe District	Safer District	Less Safe District	Safer District	Less Safe District
CALCULATIONS BASED ON ALL RESPONDENTS WITH NON-MISSING DATA								
1978	.676 (695)*	.635 (731)	69.2 (754)	67.7 (794)	.648 (994)	.605 (963)	1.21 (994)**	1.10 (963)
1980	.914 (372)**	.848 (472)	66.1 (462)*	63.7 (589)	.705 (569)	.603 (713)	1.28 (569)**	1.14 (713)
1982	.882 (305)	.850 (428)	65.8 (411)	63.9 (574)	.648 (563)	.664 (681)	1.29 (544)***	1.04 (670)
1984	.881 (495)	.880 (860)	61.8 (533)	62.6 (892)	.572 (670)	.665 (1112)	1.22 (670)**	1.11 (1112)
1986	.869 (505)	.886 (815)	64.5 (600)	67.0 (911)	.589 (791)	.719 (1155)	1.27 (791)*	1.20 (1155)
1988	.876 (522)	.933 (672)	63.0 (632)	65.8 (758)	.609 (755)	.753 (892)	1.25 (755)*	1.18 (892)
1990	.840 (676)	.881 (631)	64.6 (713)	67.7 (693)	.597 (921)	.643 (841)	1.26 (921)	1.25 (841)
1992	.837 (461)	.836 (692)	60.2 (560)	61.5 (789)	.400 (877)	.515 (1059)	1.31 (864)***	1.15 (1051)
1994	.867 (506)***	.775 (592)	62.1 (590)*	59.2 (662)	.531 (724)	.482 (778)	1.26 (724)***	1.06 (778)
1996	.835 (534)	.806 (503)	60.5 (569)	61.5 (579)	.537 (762)	.542 (675)	1.15 (762)**	1.04 (675)
1998	.821 (427)	.843 (392)	60.0 (423)	61.7 (415)	.456 (628)	.549 (509)	1.14 (628)***	.95 (509)
2000	.881 (368)*	.833 (570)	62.7 (442)	61.7 (621)	.447 (713)	.534 (874)	1.21 (714)***	1.01 (874)
2002	- - -	- - -	59.1 (485)	60.7 (516)	- - -	- - -	1.25 (568)***	1.02 (595)
2004	.874 (260)	.849 (378)	64.9 (301)	65.2 (418)	- - -	- - -	1.38 (382)***	1.06 (503)
Pooled Cross-Sectional Results	.841 (6126)	.835 (7737)	63.3 (7475)	63.8 (9211)	.560 (8967)	.613 (10252)	1.25 (9886)***	1.10 (11331)

CALCULATIONS BASED ON HOUSE VOTERS ONLY

1978	.735 (373)	.697 (406)	71.8 (382)	70.5 (427)	.988 (420)	.850 (459)	1.27 (420)**	1.09 (459)
1980	.926 (283)**	.863 (343)	68.4 (326)*	65.4 (406)	.966 (352)*	.776 (450)	1.34 (352)***	1.09 (450)
1982	.901 (202)*	.829 (269)	69.4 (250)*	66.1 (326)	1.004 (276)	.868 (342)	1.44 (274)***	1.04 (342)
1984	.894 (322)	.889 (595)	63.8 (348)	65.1 (615)	.805 (385)	.918 (680)	1.29 (385)**	1.11 (680)
1986	.879 (306)	.891 (439)	67.4 (335)	70.8 (468)	.914 (374)	1.127 (513)	1.33 (374)	1.26 (513)
1988	.899 (388)	.949 (449)	66.4 (435)	68.8 (473)	.874 (469)	1.082 (513)	1.34 (469)*	1.22 (513)
1990	.874 (333)	.904 (313)	69.5 (339)	70.6 (329)	1.045 (374)	1.023 (349)	1.42 (374)	1.33 (349)
1992	.852 (356)	.857 (516)	61.8 (392)	63.5 (566)	.642 (483)	.765 (625)	1.41 (483)***	1.15 (624)
1994	.879 (314)***	.782 (382)	65.0 (340)***	59.2 (404)	.847 (360)	.754 (430)	1.40 (360)***	1.02 (430)
1996	.857 (408)	.815 (385)	62.5 (380)	62.0 (394)	.795 (456)	.752 (431)	1.24 (456)**	1.07 (431)
1998	.884 (189)	.847 (222)	64.7 (192)	65.7 (232)	.909 (223)	.788 (250)	1.28 (223)***	.98 (250)
2000	.922 (238)***	.843 (398)	67.7 (267)**	63.4 (425)	.984 (294)	.844 (462)	1.43 (294)***	1.02 (462)
2002	- - -	- - -	64.9 (240)	63.4 (303)	- - -	- - -	1.35 (248)***	1.01 (310)
2004	.900 (203)	.863 (272)	67.3 (229)	67.3 (308)	- - -	- - -	1.49 (250)***	1.14 (330)
Pooled Cross-Sectional Results	.872 (3916)**	.851 (4989)	66.4 (4457)	65.8 (5675)	.885 (4466)	.884 (5505)	1.35 (4961)***	1.12 (6144)

Notes: For each election year, cutting point between safer and less safe districts is determined by mean value of district partisan homogeneity computed across all respondents with incumbent running for reelection. Entries in parentheses are numbers of cases. Significant inter-district differences are indicated by asterisks following entries in Safer Districts column.
***significant at .001 level (one-tail t-test); **significant at .01 level (one-tail t-test); *significant at .05 level (one-tail t-test).
Analysis in 1992, 1994, 1996, 1998, 2000, 2002, and 2004 has been weighted, as well as that applying to Pooled Cross-Sectional Results.

identifiers indeed become less favorable as district partisan homogeneity rises. Second, I verify whether constituents in fact detect the ideological extremity of safer seat incumbents. Finally, I ascertain whether perceptions of member extremity lead opposition party identifiers to hold less friendly evaluations. In order to make presentation of the results more manageable, analysis in each of the three steps will only be based upon pooled cross-sectional data.

The first step simply involves, in the case of thermometer ratings and the net positivity measure, regression of the evaluations on the district partisan homogeneity and respondent partisanship variables defined above, and the interaction between these two variables. In addition, the dummy variables needed to delineate election year fixed effects are included, with 1978 as the omitted category. When incumbent job approval is the dependent variable, probit analysis is used instead. If increasing values of district partisan homogeneity lead respondents of the opposition party to have more negative opinions of the member, the coefficient of the homogeneity variable will be negative (i.e., this coefficient by itself represents the effect of homogeneity when respondent partisanship equals 0, the value for identifiers of the opposition party, because the interaction term then will likewise have a value of 0). I would also expect positive coefficients to exist for the interaction term, signifying that as values of partisanship move toward the value of 2 used for incumbent party identifiers, homogeneity does not have the same deleterious effect on evaluations that it does for opposition party identifiers.

The outcome of this analysis appears in Table 2. My expectations, it will be seen, basically are fulfilled. With all respondents included in the analysis, each of the three partisan homogeneity coefficients is significantly negative, and each of the three coefficients for the interaction term is significantly positive. Limiting the analysis to just voters leaves each sign of these six coefficients as hypothesized, even though the partisan homogeneity and interaction coefficients in the job approval equation are now insignificant, as is the interaction coefficient in the net positivity analysis. Taking all equations into account, therefore, the general finding is that identifiers with the opposition party do fall off in support for their representative as their district becomes more lopsided for the incumbent party. At the other extreme, incumbent party identifiers fall off on thermometer ratings and net positivity at a slower pace as homogeneity rises, or, in the case of both job approval analyses, modestly increase in support with greater homogeneity (i.e., in contrast to the analyses using thermometer ratings and net positivity, the sum of the homogeneity coefficient plus two times the interaction coefficient is positive when incumbent approval is used).

Table 2. The Effects of District Partisan Homogeneity on Evaluations of House Incumbents

Variable	All Respondents: Job Approval Equation	All Respondents: Thermometer Ratings Equation	All Respondents: Net Positivity Equation	Voters Only: Job Approval Equation	Voters Only: Thermometer Ratings Equation	Voters Only: Net Positivity Equation
Partisanship	.350***(.016)	6.204***(.204)	.288***(.012)	.470***(.022)	7.732***(.249)	.415***(.019)
District partisan homogeneity	-.553** (.205)	-18.835***(2.793)	-.826***(.164)	-.100 (.267)	-10.159** (3.656)	-.740** (.276)
Partisanship*District partisan homogeneity	.316* (.142)	7.871***(1.752)	.245** (.104)	.112 (.189)	3.814* (2.225)	.097 (.168)
1980	.774***(.067)	-3.858***(.838)	.014 (.048)	.700***(.085)	-4.552***(1.038)	-.057 (.074)
1982	.746***(.070)	-3.433***(.857)	.049 (.049)	.550***(.090)	-3.846***(1.110)	.003 (.080)
1984	.822***(.057)	-5.867***(.770)	.015 (.044)	.715***(.075)	-6.323***(.971)	-.024 (.070)
1986	.797***(.058)	-2.711***(.759)	.031 (.043)	.642***(.080)	-2.498* (1.015)	.092 (.073)
1988	.963***(.063)	-4.157***(.775)	.055 (.045)	.925***(.084)	-4.163***(.985)	.040 (.071)
1990	.674***(.056)	-2.858***(.772)	-.030 (.044)	.623***(.084)	-2.520* (1.064)	.044 (.077)
1992	.581***(.057)	-8.020***(.782)	-.181***(.043)	.499***(.073)	-8.931***(.972)	-.242***(.069)
1994	.534***(.057)	-7.769***(.797)	-.111* (.046)	.411***(.077)	-9.379***(1.034)	-.115 (.075)
1996	.588***(.059)	-6.974***(.817)	-.056 (.046)	.484***(.075)	-8.570***(1.024)	-.120 (.073)
1998	.642***(.064)	-6.962***(.899)	-.091 (.050)	.627***(.095)	-5.593***(1.220)	-.039 (.087)
2000	.703***(.062)	-6.044***(.835)	-.102* (.045)	.645***(.083)	-6.202***(1.054)	-.011 (.076)
2002	—	-8.022***(.855)	—	—	-6.716***(1.134)	—
2004	.699***(.073)	-3.389***(.949)	—	.610***(.094)	-4.553***(1.136)	—

. . . table continues

Table 2. The Effects of District Partisan Homogeneity on Evaluations of House Incumbents (continued)

Variable	All Respondents			Voters Only		
	Job Approval Equation	Thermometer Ratings Equation	Net Positivity Equation	Job Approval Equation	Thermometer Ratings Equation	Net Positivity Equation
Constant	.012 (.039)	61.586***(.583)	.318***(.033)	.054 (.053)	62.225***(.772)	.456***(.056)
R^2 (or pseudo-R^2)	.370	.092	.045	.381	.129	.064
N of Cases	13954	16763	19152	9013	10337	10036

Notes: Probit coefficients appear in the case of Job Approval equations, and regression coefficients in the case of Thermometer Ratings and Net Positivity equations. Entries in parentheses are standard errors of coefficients. Goodness of fit for Job Approval equations is indicated by McKelvey-Zavoina pseudo-R^2.

***Significant at .001 level; **significant at .01 level; *significant at .05 level. Two-tail t-tests apply to coefficients of Constant and election year dummy variables, and one-tail t-tests to all other coefficients.

But are opposition party identifiers, in fact, reacting against the ideological extremity of members from safer districts when they downgrade evaluations of these members? Obviously, this step of the analysis means determining whether there actually is a relationship between district partisan homogeneity and constituent perceptions of ideological extremity. In Table 3, I regress these ideological perceptions on respondent partisanship, district homogeneity, the interaction between partisanship and homogeneity, and the fixed effects election year variables. The dependent variable is defined as follows: Respondent perception of member's ideological extremity (for Republican member, score is perception of member's ideology on ANES seven point scale, where 1 is most liberal and 7 is most conservative; for Democratic member, score is reversed so that 1 is most conservative and 7 is most liberal). Thus, for Republican and Democratic incumbents alike, larger values denote placements associated with the more extreme side of their respective party's ideology.

Three election years—1984, 1988, and 1992—must be eliminated from this analysis because no questions tapping perceptions of members' ideology were asked by the ANES. Furthermore, since respondents had more difficulty in placing their member's ideology than they did in favorably or unfavorably evaluating the member, the number of respondents appearing in the Table 3 analysis is a good deal smaller than the number for the corresponding years that formed part of the Table 2 analysis.[9]

The Table 3 results plainly show that regardless of whether all respondents or just voters alone are considered, perceptions of incumbent ideology become more extreme as district partisan homogeneity rises. So constituents in more homogeneous districts indeed are able to pick up on the actual extremity of their member that has been documented in aggregate data studies through application of roll call-based indices such as DW-NOMINATE scores. In addition, respondents' partisanship influences where they place the member ideologically; i.e., opposition party identifiers see their member as more extreme than do identifiers from the member's party. The consequence, therefore, is a two-fold estrangement of opposition party identifiers from the member. Not only are they themselves more likely to be on the opposite side of the ideological scale from the member than are incumbent party identifiers, but their distance from the member is further amplified by perceptions of him or her as more extreme. On the other hand, neither interaction term in the two equations is significant, meaning that growth in perceptions of member extremity as districts become safer does not vary between opposition and incumbent party identifiers.

The final step to be taken in this three-step sequence of investigations is determining whether perceptions by constituents of greater member extremity affect the favorableness of their evaluations. Previous studies based upon

Table 3. The Effects of District Partisan Homogeneity on Perceptions of House Incumbents' Ideological Extremity

Variable	All Respondents	Voters Only
Partisanship	-.078***(.018)	-.057** (.021)
District partisan homogeneity	1.333***(.237)	1.315***(.305)
Partisanship* district partisan homogeneity	-.025 (.149)	.022 (.185)
1980	.085 (.070)	.033 (.080)
1982	.139 (.075)	.129 (.129)
1986	-.056 (.060)	-.138 (.075)
1990	-.050 (.062)	-.055 (.079)
1994	.206***(.062)	.217**(.075)
1996	.286***(.065)	.307***(.077)
1998	.214** (.068)	.259**(.086)
2000	.278***(.066)	.262***(.078)
2002	.594***(.066)	.517***(.087)
2004	.255***(.073)	.209*(.083)
Constant	4.405***(.050)	4.515***(.061)
R^2	.030	.029
N of Cases	9501	6033

Note: Entries in parentheses are standard errors of regression coefficients.
***Significant at .001 level; **significant at .01 level; *significant at .05 level. Two-tail t-tests apply to coefficients of Constant and election year dummy variables, and one-tail t-tests to all other coefficients.

aggregate data have established that an incumbent's reelection margin and probability of victory are impaired by more extreme roll call voting (Brady et al. 2000, 184-89; Ansolabehere et al. 2001, 146, 151; Canes-Wrone et al. 2002, 132-37; Erikson and Wright 2009, 85-88). But these aggregate data studies cannot, of course, untangle the relative responsibility of incumbent party versus opposition party identifiers for the phenomenon. Here, I explain each incumbent evaluation in terms of respondent partisanship, perceptions of incumbent ideological extremity, the interaction between these two variables, and the fixed effects election year variables.

Table 4 includes the coefficients from the estimation. Similar to what was uncovered in Table 2 when I focused on district partisan homogeneity as the chief independent variable, opposition party and incumbent party identifiers are now found to react differently to increasing ideological extremity. With respondent partisanship equal to 0 for opposition party identifiers, the effect of increasing extremity—indicated simply by the coefficient of the extremity variable—is significantly negative in all equations. The coefficient of the interaction term is always significantly positive, so the impact of greater extremity on evaluations becomes less negative as the values of partisanship rise. When partisanship equals 2 for incumbent party identifiers, the effect of extremity—measured by the sum of the extremity coefficient plus two times the interaction term coefficient—is always positive, signifying for the representatives of these respondents that extremity is rewarded.

One other item of interest in Table 4 is that the coefficients of respondent partisanship are consistently negative. Interpreting this seeming anomaly requires a more detailed consideration of how partisanship operates in tandem with ideological extremity. This is illustrated at the bottom of the table by determining what the impact on each evaluation would be of moving from a perception of 1 to a perception of 7 on the ideological extremity scale. In calculating each predicted evaluation score, I assume that all respondents had a fixed combination of partisanship and ideological perception (e.g., identifying with the opposition party and placing their member at 6 ideologically), but maintained their actual values on the election year dummy variables. The predicted scores show that members perceived to be at the very low end of the ideological scale—and consequently far removed from the norm for their party—are actually more highly evaluated by identifiers with the opposition party than by their own party identifiers. This, therefore, explains the negative signs of the partisanship coefficients in the equations. But by the time ideological extremity values reach 3 (or 2 in the case of the thermometer score analysis involving all respondents), incumbent party partisans view the member more positively than do opposition partisans. These differences between their evaluations subsequently widen with growing extremity, as the fall off in opposition party identifiers' evaluations accelerates much more quickly than does growth in incumbent party identifiers' evaluations.

The conclusion of the three-step analysis extending across Tables 2, 3, and 4, then, is clear. Incumbents from districts with greater homogeneity are only modestly more popular, because opposition party constituents have particularly critical evaluations of these members. Their negative evaluations seem to be a reaction to the perceived ideological extremity of members from safer districts; incumbent party identifiers, on the other hand, only modestly reward greater extremity with higher evaluations. Thus, this negativity on the part of opposition party identifiers helps to offset the simple fact of greater numbers of incumbent party identifiers who reside in safer districts.

Table 4. The Effects of Perceptions of House Incumbents' Ideological Extremity on Incumbent Evaluations

	All Respondents			Voters Only		
Variable	Job Approval Equation	Thermometer Ratings Equation	Net Positivity Equation	Job Approval Equation	Thermometer Ratings Equation	Net Positivity Equation
Partisanship	-.418***(.067)	-5.550***(.802)	-.474***(.064)	-.551***(.094)	-7.134***(1.021)	-.702***(.089)
Perception of member's ideo-logical extremity	-.293***(.020)	-5.186***(.253)	-.285***(.020)	-.382***(.026)	-6.411***(.323)	-.427***(.028)
Partisanship*Perception of member's ideo-logical extremity	.181***.014)	2.872***(.166)	.199***(.013)	.226***(.019)	3.462***(.208)	.262***(.018)
1980	.697***(.085)	-4.379***(1.080)	-.084 (.078)	.628***(.103)	-4.904***(1.235)	-.136 (.097)
1982	.658***(.091)	-3.942***(1.146)	.035 (.082)	.526***(.113)	-3.987** (1.361)	-.040 (.107)
1986	.685***(.069)	-3.306***(.925)	-.032 (.066)	.594***(.092)	-2.636* (1.147)	.027 (.090)
1990	.600***(.069)	-4.106***(.958)	-.139* (.068)	.656***(.101)	-2.857* (1.222)	-.012 (.096)
1994	.462***(.068)	-8.286***(.956)	-.285***(.068)	.381***(.088)	-8.845***(1.162)	-.209* (.091)
1996	.549***(.074)	-6.400***(.993)	-.168* (.072)	.489***(.091)	-7.183***(1.178)	-.122 (.093)
1998	.690***(.078)	-6.116***(1.061)	-.168* (.075)	.734***(.111)	-4.810***(1.335)	-.105 (.105)
2000	.626***(.076)	-6.084***(1.017)	-.044 (.073)	.658***(.097)	-5.373***(1.202)	-.003 (.095)
2002	—	-9.346***(1.054)	—	—	-7.038***(1.338)	—
2004	.640***(.085)	-4.396***(1.104)	—	.597***(.106)	-4.619***(1.275)	—
Constant	1.392***(.104)	85.120***(1.373)	1.734***(.107)	1.856***(.142)	90.877***(1.759)	2.463***(.151)
R^2 (or pseudo-R^2)	.394	.162	.095	.420	.218	.130
N of Cases	7708	8947	8034	5159	5854	5070

	Oppos. Party	Inc. Party	Oppos. Party	Inc. Party	Oppos. Party	Inc. Party	Oppos. Party	Inc. Party	Oppos. Party	Inc. Party	Oppos. Party	Inc. Party
Ideology = 1	.946	.874	74.8	69.5	1.343	.793	.974	.906	79.7	72.3	1.970	1.090
Ideology = 2	.907	.888	69.7	70.0	1.058	.906	.942	.916	73.2	72.8	1.543	1.187
Ideology = 3	.850	.900	64.5	70.6	.773	1.019	.885	.926	66.8	73.3	1.116	1.284
Ideology = 4	.774	.911	59.3	71.2	.488	1.132	.797	.936	60.4	73.9	.689	1.381
Ideology = 5	.680	.921	54.1	71.7	.203	1.245	.676	.944	54.0	74.4	.262	1.478
Ideology = 6	.571	.931	48.9	72.3	-.082	1.358	.533	.951	47.6	74.9	-.165	1.575
Ideology = 7	.457	.939	43.7	72.8	-.367	1.471	.385	.958	41.2	75.4	-.592	1.672

Notes: Probit coefficients appear in the case of Job Approval equations, and regression coefficients in the case of Thermometer Ratings and Net Positivity equations. Entries in parentheses are standard errors of coefficients. Goodness of fit for Job Approval equations is indicated by McKelvey-Zavoina pseudo-R^2. Entries at bottom are incumbent evaluation levels predicted at each combination of respondent partisanship and incumbent ideology.
***Significant at .001 level; **significant at .01 level; *significant at .05 level. Two-tail t-tests apply to coefficients of Constant and election year dummy variables, and one-tail t-tests to all other coefficients.

The Effects of Casting a Pro-Incumbent Vote on Evaluations of the Incumbent

The finding that incumbents have little more favorability in more homogeneous districts poses a serious challenge to Brunell's prescription that drafters of congressional district lines should carve out the maximum number of safe seats for each party. I have not yet directly examined, however, his argument that the act of voting for incumbents in and of itself heightens support for them. This, after all, was supposed to be the chief individual-level mechanism behind his aggregate-level inference that safer seats have higher levels of incumbent approval, in that greater amounts of pro-incumbent voting and hence more widespread approval should occur where more constituents belong to the district's majority party.

Brunell's case empirically rests upon his demonstration that in equations deploying post-election evaluations of winning incumbents as the dependent variable, the independent variable of whether the constituent voted for the incumbent is considerably significant. The direction of causality, however, cannot be established from such a demonstration. It certainly is possible that what underlies the relationship is that constituents who had positive evaluations prior to the election simply voted for the incumbent as a result of these evaluations, and that the positive evaluations persisted after the election with no intervention from the voting act itself. Thus, any test of the proposition that pro-incumbent voting is an independent determinant of more positive incumbent evaluations should go beyond Brunell's analysis to include as a right-hand-side variable a measure of how the member was judged prior to election day.

For my own individual-level analysis, I analyze the effect of voting decisions on incumbent evaluations whenever the ANES data permit the creation of a corresponding lagged evaluation variable. The 1996 study was unique in including an identical incumbent evaluation question (i.e., thermometer ratings) in both its pre- and post-election waves. All other ANES studies that ask evaluation questions do so on a post-election basis only. It is possible, however, to take advantage of the fact that some of these studies are part of a larger panel study spanning more than one election. Thus, for example, an equation using 2002 post-election thermometer ratings as the dependent variable can be devised in which respondents' lagged thermometer ratings of the same incumbent from the 2000 post-election study appear as a right-hand-side variable.

There is one other reason to revisit Brunell's analysis. The right-hand-side incumbent voting variable is clearly endogenous to the system; i.e., determined by many of the same exogenous factors that determine incumbent evaluation itself. As such, it very likely is correlated with the error term (ε) in the equation, barring the improbable event that the exogenous independent variables in the equation form the full set of all possible variables that are directly related to both incumbent evaluation and incumbent voting (Timpone 2003, 292). Failure to account for this endogeneity would produce inconsistent estimation of the incumbent voting coefficient.[10]

I employ the standard solution to such a problem by substituting for the endogenous voting variable an instrumental variable, equal to the predicted values of incumbent voting based upon its reduced form equation. In the reduced form (i.e., first stage) equation, the predictive factors to be used will be the predetermined (i.e., exogenous and lagged endogenous) variables that are direct causes of either incumbent evaluations or of incumbent voting. The instrumental variable, like the predetermined variables that have been employed in its formulation, will then itself be uncorrelated with the error term when it is substituted for incumbent voting in the second stage equation explaining incumbent evaluations (Kennedy 1998, 139-40).

The second stage equation, besides using this instrumental variable and lagged incumbent evaluations as right-hand-side variables, also includes the following set of exogeneous control variables:

Partisanship (as defined above)

Perceived economic change (for respondent with House incumbent of presidential party, 3 if U.S. economy seen as better over past year, 2 if same, 1 if worse; for respondent with incumbent of non-presidential party, coding is reversed)

Race (for respondent with Republican incumbent, 1 if respondent is white, 0 if non-white; for respondent with Democratic incumbent, coding is reversed)

Union membership (for respondent with Republican incumbent, 1 if neither respondent nor any other member of household belongs to union, 0 otherwise; for respondent with Democratic incumbent, coding is reversed)

Southern residency (for respondent with Republican incumbent, 1 if respondent resides in southern state, 0 otherwise; for respondent with Democratic incumbent, coding is reversed)[11]

Incumbent spending (natural logarithm of incumbent's campaign spending in $1000s)

Challenger spending (natural logarithm of challenger's campaign spending in $1000s)[12]

Tenure (1 if respondent's incumbent is non-freshman, 0 if freshman)[13]

Age (natural logarithm of respondent's age in years)

Education (6 if respondent has advanced degree, 5 if bachelor's degree, 4 if some college education or junior or community college degree, 3 if high school diploma, 2 if some high school education, 1 if grade school diploma or less)

Public affairs follows what's going on in government and public affairs "most of the time," 3 if "some of the time," 2 if "only now and then," 1 if "hardly at all")

Trust in government (4 if respondent thinks government in Washington can be trusted to do what's right "just about always," 3 if "most of the time," 2 if "only some of the time," 1 if "never")

Member's party (1 if respondent's member is Republican, 0 if Democrat).

The first six of these control variables should have a positive relationship with incumbent evaluations. (Southern residency, however, might be expected to have migrated toward a positive relationship across the 1990s as the South became increasingly Republican at the congressional level.) Challenger spending should negatively affect incumbent evaluations, in that much of it is aimed at castigating the member rather than building up the challenger. Consequently, a plurality of television ads run by challengers have been found to be negative attack ads (Goldstein et al. 2001, 97). Freshman status may depress evaluations simply because of the more limited time that has been available to cultivate constituent favor through casework and porkbarreling. Older respondents would have had more opportunity to learn of the emphasis that members in general, including their own, place upon constituency involvement. By the same token, better educated respondents and those with greater interest in government and public affairs would have more of the information necessary to be aware of this involvement. Generalized trust in government might well extend to the point of greater confidence in one's own officeholder. Finally, the member's party is included to account for election-to-election differences in the party benefiting from national political tides.

All the above predetermined predictors of incumbent evaluations, as previously mentioned, must also be used in the first stage reduced form equation to create the instrumental variable for incumbent voting. Furthermore, two other variables that are expected to be predictors only of incumbent voting are included in the reduced form equation:

Recognize challenger's name (1 if respondent claims to recognize House challenger's name when asked to rate challenger on thermometer scale, 0 otherwise)

Offer comment about challenger (1 if respondent is able to offer at least one open-ended comment—either positive or negative—about House challenger, 0 otherwise).[14]

Both of these variables allow demarcation of especially low-visibility challengers; regardless of feelings about the incumbent, respondents will be reluctant to vote for a challenger they have faint or no knowledge of. By way of contrast, there is little reason to think that inability to recognize or comment on the challenger would directly affect how incumbents are evaluated by the public.

Voting for an incumbent winner is a dichotomous variable (with a pro-incumbent vote coded as 1 and a vote for the challenger as 0), so the instrumental variable based upon its predicted values in the reduced form equation will be calculated using probit analysis. These predicted values are the raw untransformed predictions, rather than the transformed probabilities (Alvarez and Glasgow 1999, 150). Substituting the predicted for the actual values of incumbent voting, however, will yield biased standard errors of the coefficients in the second stage estimation, regardless of whether the second stage dependent variable is continuous (i.e., thermometer ratings or net positivity) or dichotomous (i.e., job approval). When the second stage dependent variable is continuous (here, the procedure is sometimes termed two stage probit least squares (2SPLS)), the correction is to multiply the standard errors by a weighting factor (Alvarez and Glasgow 1999, 150). But when the second stage dependent variable is dichotomous, serious problems are posed for the standard error correction (Timpone 2003, 299). Consequently, my analysis of how voting affects incumbent evaluations will only be based upon the two continuous measures, with the standard error correction applied. Taking into account the requirement stated above that lagged evaluation measures must be available for the estimation, this leaves a total of seven analyses that can be undertaken, involving thermometer ratings in 1992, 1994, 1996, and 2002, and net positivity in 1992, 1994, and 1996.[15] Because of the limited number, I shall estimate seven separate equations rather than doing pooled analysis.[16]

Tables 5 and 6 display the results of the 2SPLS analysis. In Table 5, the probit coefficients of the first stage reduced form equations appear. The most essential feature of the equations is that all seven yield robust instrumental variables for the endogenous incumbent voting variables. The McKelvey-Zavoina pseudo-R^2 statistics show summary goodness of fit ranging from .593 to .836. Alternatively, comparing how well the equations predict respondents' actual values on the dependent variable with null model predictions simply forecasting all values to be at the mode (i.e., a pro-incumbent vote), I find improvement ranging from 6.8 to 15.0 percentage points. Also important is that the two exogenous factors added to the reduced form equations that were not direct causes of incumbent evaluation—recognizing the challenger's name and being able to offer at least one open-ended comment about the challenger—do indeed function as determinants of

Table 5. First Stage Reduced Form Equations Used to Generate Instrumental Variables for Incumbent Voting

Variable	1992 Thermometer Ratings	1992 Net Positivity	1994 Thermometer Ratings	1994 Net Positivity	1996 Thermometer Ratings	1996 Net Positivity	2002 Thermometer Ratings
Lagged incumbent evaluations	.026***(.005)	.299***(.065)	.028** (.009)	.504***(.126)	.032***(.004)	.245***(.051)	.018** (.006)
Partisanship	.696***(.116)	.618***(.098)	1.104***(.204)	1.248***(.194)	.882***(.084)	.950***(.081)	.755***(.135)
Perceived economic change	.416** (.166)	.505***(.153)	.225 (.256)	.249 (.214)	.245** (.103)	.352***(.100)	-.012 (.233)
Race	-.101 (.475)	.208 (.408)	.646 (.563)	.711 (.500)	-.110 (.216)	-.272 (.205)	-.270 (.403)
Union membership	-.223 (.251)	-.074 (.223)	.831* (.399)	.625* (.363)	-.021 (.169)	.069 (.166)	.690* (.303)
Southern residency	.502 (.321)	.555* (.267)	-.124 (.407)	.452 (.381)	.138 (.173)	.030 (.164)	-.332 (.296)
Incumbent spending	.345* (.184)	.339* (.155)	.393 (.476)	.351 (.420)	.073 (.056)	.048 (.058)	.209 (.243)
Challenger spending	-.094 (.077)	-.062 (.065)	-.235* (.128)	-.222* (.114)	.071 (.037)	.064 (.035)	-.174**(.074)
Tenure	-.182 (.323)	-.369 (.276)	-.353 (.468)	.151 (.362)	.158 (.180)	.130 (.176)	.034 (.385)
Age	-.046 (.272)	.109 (.232)	.266 (.512)	.401 (.439)	.028 (.234)	.095 (.232)	-1.059 (.420)
Education	.201* (.097)	.142* (.085)	.086 (.134)	.321* (.139)	.009 (.063)	.056 (.061)	-.147 (.109)
Public affairs	-.178 (.131)	-.114 (.116)	.413* (.242)	.091 (.199)	.130 (.087)	-.014 (.081)	.023 (.164)
Trust in government	.325* (.191)	.121 (.159)	-.948 (.388)	-.801 (.312)	-.074 (.140)	.080 (.134)	.192 (.216)
Member's party	1.082 (.577)	.891 (.505)	-.340 (.671)	-.093 (.569)	.724** (.260)	.840***(.250)	-.548 (.552)
Recognize challenger's name	-.006 (.240)	.040 (.208)	-1.188* (.526)	-.884* (.411)	-.241 (.209)	-.247 (.194)	-.991** (.397)
Offer comment about challenger	-.815** (.270)	-.567* (.250)	-.984** (.400)	-1.002** (.369)	-.637***(.158)	-.565***(.154)	—

Constant	-5.284** (1.829)	-4.289** (1.479)	-3.082 (3.219)	-3.469 (2.823)	-3.553** (1.137)	-2.320* (1.123)	-3.537 (2.539)
Pseudo-R^2	.634	.593	.825	.836	.714	.660	.675
N of cases	341	397	200	234	758	705	245

Note: The dependent variable for all equations is incumbent voting. Entries in parentheses are standard errors of probit coefficients. Analyses have been weighted. Goodness of fit is indicated by McKelvey-Zavoina pseudo-R^2.
***Significant at .001 level; **significant at .01 level; *significant at .05 level. Two-tail t-tests apply to coefficients of Incumbent's Party and Constant, and one-tail t-tests to all other coefficients.

Table 6. Second Stage Estimation of Incumbent Evaluations

Variable	1992 Thermometer Ratings	1992 Net Positivity	1994 Thermometer Ratings	1994 Net Positivity	1996 Thermometer Ratings	1996 Net Positivity	2002 Thermometer Ratings
Lagged incumbent evaluations	.404***(.091)	.312** (.118)	.449***(.076)	.317** (.095)	.554***(.052)	.330***(.061)	.427***(.067)
Instrumental variable for incumbent voting	4.149 (2.984)	.366 (.354)	2.973* (1.648)	.332* (.144)	5.094***(1.378)	.552** (.189)	3.478 (2.298)
Partisanship	.781 (2.327)	-.024 (.231)	1.770 (2.410)	-.073 (.213)	-.403 (1.373)	-.164 (.197)	2.738 (2.133)
Perceived economy change	1.022 (2.115)	-.007 (.227)	-.545 (1.827)	-.150 (.138)	.302 (.709)	-.053 (.102)	-3.094 (1.829)
Race	5.765* (3.501)	-.315 (.292)	.977 (4.135)	-.498 (.321)	1.609 (1.427)	.224 (.188)	2.433 (3.228)
Union membership	-.285 (2.307)	-.031 (.181)	-.940 (3.468)	-.118 (.254)	.535 (1.084)	-.096 (.141)	3.589 (3.013)
Southern residency	1.030 (3.276)	-.101 (.308)	-.501 (2.842)	-.597 (.216)	-1.247 (1.067)	.044 (.130)	3.823* (2.311)
Incumbent spending	.350 (1.794)	.002 (.169)	-5.587 (3.040)	-.305 (.233)	.804* (.381)	-.025 (.048)	-1.458 (2.072)
Challenger spending	-1.092 (.789)	-.012 (.061)	-.190 (.992)	.074 (.076)	.013 (.225)	.024 (.028)	.077 (.747)
Tenure	2.291 (2.992)	.393 (.262)	2.661 (3.005)	.514* (.226)	-.194 (1.202)	.174 (.149)	-4.595 (3.092)
Age	5.732* (2.641)	.541** (.204)	4.677 (4.000)	.599* (.309)	3.022* (1.406)	.694*** (.181)	6.515* (3.914)
Education	.663 (.918)	.088* (.073)	-1.179 (1.077)	-.099 (.092)	-.125 (.389)	.062 (.049)	.390 (.946)
Public affairs	-.427 (1.411)	.195* (.110)	3.289* (1.587)	.341** (.115)	-.297 (.532)	.107 (.068)	-1.242 (1.375)
Trust in government	3.316* (2.015)	-.206 (.140)	.931 (3.132)	-.061 (.227)	1.335 (.888)	.002 (.112)	2.995* (1.710)
Member's party	-5.575 (5.901)	.015 (.502)	-5.769 (4.600)	-.018 (.347)	-1.251 (1.849)	-.085 (.260)	-7.441 (4.964)

Constant	2.141 (22.088)	-2.025 (1.872)	41.866 (22.959)	-.324 (1.737)	4.105 (8.545)	-3.101**(1.003)	18.850 (21.676)
R^2	.461	.308	.507	.390	.686	.332	.537
N of cases	341	397	200	234	758	705	245

Note: The dependent variable for each equation is the incumbent evaluation named at the top of the column. Entries in parentheses are standard errors of regression coefficients. Analyses have been weighted.

***Significant at .001 level; **significant at .01 level; *significant at .05 level. Two-tail t-tests apply to coefficients of Incumbent's Party and Constant, and one-tail t-tests to all other coefficients.

the vote. This is a statistical requisite in order for the second stage equations to be identified (Timpone 2003, 303). The former variable achieves significance in the expected negative direction three times, as does the latter variable every time that it appears.

The second stage equations with incumbent evaluations as the dependent variable are presented in Table 6. Not surprisingly, the control variables here are less potent predictors than they were in the reduced form equations, given that they must compete for explanatory relevance with the instrumental variables that are now part of the equations. Fewer control variables are significant, and there are more instances where the parameters are wrongly signed. Lagged evaluations, of course, are always significant at least at the $p < .01$ level (one-tail t-test). But the instrumental variables for incumbent voting, which are at the heart of my theoretical focus, exhibit no more than a mixed record of explanatory success. A pro-incumbent vote always leads to higher incumbent evaluations, but the relationship is significant at the $p < .05$ level only in the narrowest majority of the seven equations (both 1994 and both 1996 equations).

Brunell's assessment of the effect of pro-incumbent voting on incumbent evaluations, therefore, likely overstated this effect by not accounting for either pre-election evaluations or the endogeneity of the voting variable. I must conclude from my own analysis that the act of voting for a victorious House member is only an imperfect mechanism for improving his or her standing in the voter's mind. Any prescription, like Brunell's, built upon the premise that carving out more safe districts will garner greater public esteem for members through the workings of the electoral process seems destined frequently to fall short of meeting its goals.

Summary and Conclusions

The centerpiece of what I have found in this study is that the degree of partisan homogeneity in House districts makes little difference on how positively incumbents are viewed. This is true regardless of whether all respondents with non-missing ANES data are considered or just those who voted for Congress. Part of the reason for this non-intuitive result is that opposition party identifiers in safer districts have disproportionately critical evaluations of the member that largely offset the sheer weight of the greater numbers of incumbent party identifiers residing there. These opposition party identifiers seem to be reacting against the ideological extremism of members from safer districts, which they are able to discern in their perceptions of where the member stands ideologically.

The other part of the answer to why safer and less safe districts do not differ more in terms of how favorably their members are evaluated lies in the inconsistent tendency of pro-incumbent voting to generate more positive images of the incumbent in constituents' minds. Greater numbers of votes are cast for winning incumbents in more homogeneous districts, but evidence that these greater numbers of votes translate into an elevated level of post-election incumbent esteem is present in some elections but not others. And even when some enhancement in the member's standing does follow in the wake of voting for him or her, it is possible that such an immediate boost registered in the post-election ANES may ultimately prove ephemeral, dissipating soon into the next year.

Thus, in lieu of a stronger relationship between voting and evaluations, the proposal for making districts more homogeneous basically is reduced to the argument that increasing the proportion of incumbent party identifiers—regardless of whether they vote—will simply mean higher aggregate ratings of the incumbent. But as I have shown, this conceptualization is insufficiently nuanced. Other factors, in fact, can also change as a result of such packing, such as the member being pressured to shift in a more extreme direction toward the issue positions preferred by the party base, despite strongly alienating opposition party supporters.

This means, of course, that there might as well not be any improvement in the quality of overall policy representation afforded by incumbents in districts made safer by redistricting. Brunell has contended on intuitive grounds that more homogeneous districts allow for more accurate representation of constituents' policy views, because the member will have a better sense of district opinion (2008, 14, 77-79). Yet in the most far-reaching empirical study of its kind, Griffin finds from 1972 to 2000 that the relationship between House members' roll call liberalism and the liberalism of their districts actually was greater in constituencies with a more even balance of partisanship (2006, 915-18). Likewise, on a within-district basis, individual members from more competitive districts were more likely to adjust their roll call liberalism to changes in district ideology over time (2006, 918-19). So overall policy congruence between members and constituents, which Brunell strongly acknowledges to be an integral component of constituents' contentment with their member (2008, 79), is in fact weaker in safer districts, thus providing further evidence that making districts more homogeneous is not likely to make for more satisfied constituents.

NOTES

[1]Stonecash, Brewer, and Mariani (2003, 17-129) have presented what is probably the most comprehensive investigation of the trend toward greater district partisan homogeneity.

[2]In his effort to formulate a comprehensive explanation for the growth in roll call polarization in the House, Theriault (2008, 62-108) supplies evidence to support the conclusion that all three theories of growing district partisan homogeneity have validity.

[3]Work done by Ansolabehere et al. (2006) poses a challenge to Brunell's optimism concerning the effects of primaries, in that they find low levels of competition to have characterized primaries since their inception in the early 1900s and that the overall trend has been in a downward direction. Thus, movement toward lower primary competition accompanied the much more familiar movement toward lower competition in general elections. Brunell, however, believes that given incumbents' generally intense desire to retain their jobs, only a small number of renomination defeats would suffice to keep them responsive to their constituents (2008, 12).

[4]In the case of the two midterm election years immediately following redistricting—1982 and 2002—the returns that are used to tap district partisan homogeneity are those from the preceding presidential election, recomputed within the new district lines.

[5]These cutting points are as follows: 1978 (.049), 1980 (.063), 1982 (.061), 1984 (.071), 1986 (.074), 1988 (.069), 1990 (.061), 1992 (.050), 1994 (.070), 1996 (.071), 1998 (.068), 2000 (.075), 2002 (.085), and 2004 (.100).

[6]In 1978, however, respondents were asked to rate their member's job performance on a five-point, rather than four-point, scale: very good, good, fair, poor, and very poor. Hence, respondents in the first two categories that year were recoded as approving of the incumbent, and those in the last three categories were recoded as disapproving. As a result of classifying the "fair" response as part of the latter category, the mean job approval ratings I calculate in 1978 will be somewhat lower than those calculated in the other election years when respondents were asked directly whether they approved or disapproved of job performance.

[7]Up to four positive and four negative comments were coded by the ANES from 1978 through 1986, and five comments of each type from 1988 onward. Substantively, however, this expansion in the variable's possible range of values makes little difference, given the very small numbers of cases where either five positive or five negative comments are made.

[8]Brunell has attempted to rebut the claim that the more lopsided districts he favors would lead to greater ideological polarization by showing that roll call voting does not become more extreme as the incumbent's electoral margin grows (2006, 81-82; 2008, 95-97). Mann and Erikson and Wright, however, demonstrate that when district safety is measured more appropriately in terms of underlying partisanship, greater homogeneity does in fact produce more roll call extremity. As Canon points out, reelection margin taps voters' degree of satisfaction with the incumbent, and one reason why members from non-homogeneous districts may personally be electorally secure is precisely because they moderate their roll call record (2009, 367-68).

[9]42.4 percent of all respondents across the relevant surveys from 1978 to 2004 could not answer the ideological question.

[10]Timpone also points out that such inconsistency will result regardless of whether the coefficient of the endogenous variable is being estimated as part of a single equation, or whether a system of reciprocal equations is involved.

[11]Southern states are the 11 of the old Confederacy, plus Kentucky and Oklahoma.

[12]The challenger's actual spending is always supplemented by $1 to avoid calculating the natural logarithm of zero when no spending is reported.

[13]Except for the analysis of thermometer ratings in 1996, where it was possible to rely upon the ANES pre-election wave for the lagged evaluation variable, all other analyses must exclude freshman incumbents winning their seat in a special election.

Obviously, no lagged evaluations of the incumbent from two years previously would be available.

[14]Since respondents in the 2002 ANES were not afforded the opportunity to offer open-ended comments about either House candidate, this exogenous variable cannot be employed in the creation of the instrumental variable for incumbent voting that year.

[15]Note that the analysis of incumbents' net positivity in 1996 requires use of a lagged positivity variable from 1994; i.e., unlike the case with thermometer scores in 1996, there were no open-ended evaluation questions asked about the incumbent in the pre-election wave of the 1996 study. The panel study that is used in this case is the 1992-1997 ANES Combined File. The same panel study provides the 1992 lagged evaluations for the 1994 analyses employing thermometer scores and net positivity, respectively, as the dependent variables. For the 1992 analyses using these two evaluations as the dependent variables, lagged 1990 evaluations are taken from the 1990-1992 ANES Full Panel File. Finally, the 2002 analysis using thermometer scores as the dependent variable extracts its 2000 lagged evaluations from the 2000–2002–2004 ANES Full Panel File.

[16]Because of the nationwide redistricting that preceded the 1992 and 2002 elections, voters in those years whose congressional district had changed were excluded from the analysis. Their lagged evaluations from two years earlier obviously would not have been of the incumbent candidate who was on their ballot in 1992 or 2002.

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