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Split-Ticket Voting: The Effects of Cognitive Madisonianism

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Split-ticket voting has recently received special attention, because it provides a possible microlevel explanation for institutionally divided government. Are split-ticket voters intentional, selecting one party for president and another for Congress, in order to somehow check and balance government? A general model of split-ticket voting is specified, taking into account the important but neglected interaction effects of party, candidate quality, and incumbency. Then, cognitive Madisonian variables are incorporated and logistic regression models estimated on 1992 and 1996 national election data. Strong cognitive Madisonian effects are found. Model Madisonians, who seek to divide power and balance policy, make up over 20% of the electorate and may be largely responsible for the observed patterns of division at the aggregate level.

Why do American voters split the ticket, choosing one party for the White House and another for Congress? Are they “cognitive Madisonians” (Ladd 1990), splitting the ticket after a conscious decision that it is somehow “good” to check power and balance policy, as our nation’s Founders might have wanted? Indirect evidence of opinion surveys suggests they might be. Nationwide polls in the post-World War II period have repeatedly demonstrated that Americans favor divided over unified government (Sigelman, Wahlbeck, and Buell, 1997, 880). If most Americans *want* a partisan division of government power, then that intention, one might suppose, would weigh heavily in a decision to ticket-split. Still, many relevant studies show little, if any, effect from cognitive Madisonianism. Thus, a paradox suggests itself. Voters say they prefer to check government power but do not act on that preference. The aggregate patterns of divided government, real and decisive as they are, lack that obvious microfoundation. Herein we resolve this paradox showing that American voters, at least in 1992 and 1996, clearly exercised their preference for divided government. Confirming this source of ticket-splitting seems especially important, given how many election outcomes have been decided at the edges of the system, beyond the usual reach of partisanship.

The body of the article has six sections. First, the literature, and our place in it, is discussed. Second, the items and context of the 1992 and 1996 election surveys are reviewed. Third, a baseline model of ticket-splitting, stressing the special role of party identification and candidate likeability, is developed. Fourth, a full model of ticket-splitting is estimated. Fifth, the magnitude of cognitive Madisonian effects on vote choice is assessed. Sixth, conclusions are drawn about the general importance of cognitive Madisonianism for the American voter.

Aggregate v. Individual, Separable v. Nonseparable, Balancing v. Splitting: Directions from the Literature

Published work on divided government has become voluminous. Much, but not all, is directly related to the dependent variable we wish to explain—the behavior of ticket-splitting by individual American voters. The literature can be organized into studies of aggregate versus individual data, separable versus nonseparable preferences, and policy balancing versus actual ticket splitting. The first set of investigations emphasize evidence from political geographic units—the nation or legislative districts (Burden and Kimball 1998; Fiorina 1992; Grofman et al. 2000; Jacobson 1990). These aggregate studies, while insightful, are indirect, not looking at the individual voters themselves. The second set of studies, on election survey data, do look directly at individuals. Traditionally, that work has assumed, at least tacitly, that voter preferences are separable. As Smith et al. put it, “Separable preferences can . . . be viewed simply as ‘unconditional’ desires about partisan control” (1999, 743). When statistical models with this bent are estimated, the cognitive Madisonian idea receives no support (Alvarez and Schousen 1993; Sigelman et al. 1997). Newer work, however, suggests these null results may come from the failure of the standard models to incorporate nonseparable preferences. For example, party response in a congressional election might depend on the party outcome of the presidential contest (Lacy and Paolino 1998, 22). These results are important, clearly showing that the vote equation must specify *conditional* partisan effects, otherwise they will err (see also Smith et al. 1999, 756–57). In their examination of conditional preferences, Layman, Carsey, and Rundquist (2001) are mostly concerned with explaining policy balancing, which suggests a third set of studies. In these, ticket-splitting per se is given secondary treatment, and the idea that voters split because they wish to check government power is dismissed because of measurement issues. From an extensive recent investigation of policy balancing using the NES surveys from 1976 to 1996, Mebane (2000) returns the emphasis to the dependent variable of ticket-splitting. But he finds the impact is at the margins: “a small but significant proportion of voters have been motivated to vote a split ticket in order to increase the chances of institutional balance” (51).

In sum, what does the extant literature indicate about the effects of cognitive Madisonianism on ticket-splitting in United States elections? Recalling the general notion from the Founders, that voters aim to apply “checks and balances”

to American political institutions, we apply the summary to two complementary hypotheses: checking power and balancing policy. With respect to the former hypothesis, findings are null. Specifically, respondents who feel “it is good” to divide government power have not been found significantly more likely to ticket-split. With respect to the latter hypothesis, there does appear to be significant policy balancing, variously measured. However, the effect of this policy balancing on actual ticket-splitting is marginal. Overall, then, cognitive Madisonian effects on the act of splitting the ticket appear “none to slim.” Below, we attempt to show that these conclusions do not necessarily hold.

The 1992 and 1996 Election Surveys: Items and Context

A major obstacle to testing microtheories of divided government has been the lack of relevant items in national election surveys. With respect to the American National Election Studies (ANES), only the 1992 instrument has a direct item measuring the voter’s desire to check government power by dividing it among the parties: “Is it better when one party controls both the presidency and Congress, better when control is split between the Democrats and Republicans, or doesn’t it matter?”¹ This so-called separable preferences item is employed in our central tests. A strong case for the ANES item comes initially from direct comparison to nonseparable preferences items. Here is the wording of the items Smith et al. used to determine nonseparable preferences. “If Bob Dole were to be elected president which would you prefer: a Republican Congress to help him pass his agenda or a Democratic Congress to serve as a check on his agenda?” (1999, 747). Followed by, “If Bill Clinton were to be elected president which would you prefer: a Democratic Congress to help him pass his agenda, or a Republican Congress to serve as a check on his agenda?”

For a number of reasons, the ANES item seems favored. First, its straightforwardness gives it considerable face validity. Second, it is general rather than context-specific, avoiding bias from candidate priming. (Consider the nonseparable formulation. If I profoundly dislike Bill Clinton, the question will push me to express a preference for divided government—to deny Clinton control—even though I am inclined to unified government). Third, it is a “simple” rather a “mediated” question, with far fewer “political objects” for the respondent to evaluate, and thus less potential for partisan bias. (See the original discussion in Fiorina (1981, 121–22). In the ANES item, there are five political objects (party, presidency, Congress, Democrats, Republicans), while with the nonseparable preferences item set, there are seven, five of which are repeated twice, for a total of twelve (Dole, president, Republican, Congress, agenda, Democratic, Clinton). Fourth, as linguistic statements, the ANES questions are measurably more com-

¹ The data for the 1992 ANES survey are available from the Inter-University Consortium for Political and Social Research (ICPSR), University of Michigan, Ann Arbor. Neither the ICPSR nor the gatherers of the data bear any responsibility for the interpretations offered herein.

prehensible, according to the components of the Gunning Fog Index, which assesses the educational grade level required to grasp the meaning of sentences.² Specifically, the ANES item contains 26 words, the nonseparable preferences item set 66 words. Further, the ANES item has only three “big words” (i.e., three syllables or more), whereas the nonseparable preferences item set has 12.

On the basis of these arguments, it would seem premature to dismiss the ANES item, especially if it can be found to have been replicated in other national elections surveys, beyond 1992. Fortunately, the Social Science Research Laboratory (SSRL), University of Mississippi, in their 1996 national election study, asked a highly similar question: “Is it good to vote for some candidates for each party because it helps ensure that no single party has all the power?”³ (It has four political objects, 23 words, and only one “big word” of three syllables). Further, it is like the ANES question in that it expressly asks voters about their desire to check power. What do these two items tell us about preferences for divided government in these two elections and the simple relationship of those preferences to ticket-splitting? If they indicate that few Americans want divided government, or that divided government preferences are not related to act of splitting the ticket, then we will be discouraged, not expecting more sophisticated multivariate analysis to yield much. But such is not the case.

In 1992, only 36% of the respondents believed in one-party control, a comparable number to the 1996 figure of 42% (“disagree” + “strongly disagree”). Most Americans, in other words, actively favored the checking of party power, or at least did not wish to prevent it. Furthermore, these preferences clearly relate to reported ticket-splitting behavior. In 1992, 25% of those who favored divided government said they ticket-split, in contrast to 16% among those who favored unified government. In 1996, among those who “strongly agree” government should be divided, 35% split the ticket, in contrast to only 13% among those who “strongly disagree.” For both years, these relationships are easily statistically significant, according to chi-square tests. These simple results lead us to pursue the hypothesis that cognitive Madisonianism is operating in both periods.

² The Gunning Fox Index (G) is sometimes used to assess the degree of education required to understand text. The components are sentence length and word difficulty. $G = .4(S + W)$, where S = average words per sentence and W = the percent of the words that are three syllables or more, excluding usual prefixes and suffixes. The G score suggests the years of school necessary to comprehend the text. As an example, the *International Journal of Forecasting*, an interdisciplinary scholarly statistical journal, asks authors to keep the G index of their papers below the number 16 (See “Instructions to Authors,” in a current issue of that journal).

³ The 1996 election data come from a telephone survey carried out by the Social Science Research Laboratory, University of Mississippi, October 11–November 3. The N = 995, from the District of Columbia and the states of the lower 48. Charles E. Smith and his colleagues at the University of Mississippi graciously provided us with the data set. Of course we bear full responsibility for any errors of interpretation.

TABLE 1
Partisanship and Patterns of Ticket-Splitting, 1992 and 1996

1992			
Vote Patterns	Partisan Identification (including leaners)		
	Democrat	Republican	Independent
DD	456 (79%)	29 (7%)	27 (43%)
DR	71 (12%)	21 (5%)	14 (22%)
RD	32 (6%)	86 (20%)	13 (21%)
RR	16 (3%)	300 (69%)	9 (14%)
<i>N</i> = 1,074	100%	100%	100%
	575	436	63
1996			
Vote Patterns	Partisan Identification (including leaners)		
	Democrat	Republican	Independent
DD	221 (74%)	19 (7%)	6 (21%)
DR	56 (19%)	21 (7%)	10 (35%)
RD	6 (2%)	26 (9%)	3 (10%)
RR	16 (5%)	216 (77%)	10 (35%)
<i>N</i> = 610	100%	100%	100%
	299	282	29

Ticket-Splitting and Partisanship Patterns: A Baseline Model

Ticket-splitters, having to decide in the presidential and the House contests, voted Democratic in one and Republican in the other. For the 1992 data, ticket-splitters = 22%, overall *N* = 1,074. The share of splitters for this particular election year is comparable to immediately prior contests, e.g., 24% in 1984, 25% in 1988. For the 1996 data, ticket-splitters = 20%, overall *N* = 610. In both elections, party identification strongly influences splitting behavior, as Table 1 indicates. Most obviously, identifiers are much more likely to vote a straight ticket than independents. Less obviously, party identification shapes the type of split. The first type of split—Democrat for President and Republican for the House (DR)—is committed more by Democratic partisans. The second type of split—Republican for President and Democrat for the House (RD)—is committed more

by Republican partisans. Partisan splitters, then, tend to stay with their presidential candidate, confining defection to the House race. This asymmetric splitting pattern speaks to the bigger appeal of the party's presidential candidate, relative to the more limited popularity of its own local candidate.

These tabular results suggest that party identification shapes the likelihood of ticket-splitting, but in conjunction with influences from national and local candidate appeal. Clearly, very popular candidates can cause voters to bolt party ranks. New candidates may appeal to the voters of the opposition by the programs they promise, while incumbents may carry away this opposition because of their record in office. Hence, measures of candidate attractiveness must be taken into account, along with party identification itself, to give a full accounting of ticket-splitting. This is done in the logistic regression of column 1, Tables 2a and 2b, where split-ticket voting in 1992 and 1996 is predicted from party identification (its presence and its strength), plus net presidential candidate likeability (e.g., Bush versus Clinton), net congressional candidate likeability (Republican candidate versus Democratic candidate), and whether a congressional incumbent is running. Party identification tends to decrease ticket-splitting, as expected. However, the effects of the other variables are neither strong nor systematic, their signs are difficult to interpret, and these signs tend to change from one election to another. Further, the statistical fit of the model is disappointing. The essential reason for these difficulties is misspecification, namely the impact of party identification should be conditioned by candidate appeal, rather than just entered in additive fashion, as discussed below.

Obviously, an attractive candidate can lure a voter away from his or her partisan moorings. For example, a Democrat in 1996 who liked Dole more than Clinton would more probably defect in the presidential balloting, even if he or she stayed with the party at the congressional level. That is, the impact of party identification is conditioned by presidential candidate likeability and is more properly modeled multiplicatively (i.e., identification \times likeability). Similarly with the appeal of a local congressional candidate. For instance, a Democrat in 1996 who was partial to the Republican congressional candidate was more likely to defect to that candidate, especially an incumbent one. Again, the effect of party identification depends in part on candidate appeal, and so requires specification of interaction terms. Thus, we propose a revised, basic split-ticket voting model incorporating five independent variables: strength of party identification, party identification, and party identification interacted with presidential candidate evaluation (PID \times Pres), House candidate evaluation (PID \times House), and House candidate incumbency status (PID \times INC). The logistic regression estimates of that model are reported in column 2, Tables 2a and 2b.

This additive-nonadditive model offers encouraging findings. The pattern of substantive results is essentially consistent across the two elections. The main effects of party attachment are as expected. Strong partisans are less likely to split their ticket than weak partisans or Independents. The interaction effects are where things get interesting. For instance, Democrats who felt Bush had more

TABLE 2A
Models of Split-Ticket Voting and Cognitive Madisonian Effects, 1992

	(1)	(2)	(3)	(4)	(5)
Constant	-.34 (.20)	-.13 (.18)	-.55 (.24)	.14 (.23)	-.25 (.28)
Party ID	-.09 (.11)	.33 (.13)*	.36 (.14)**	.36 (.15)**	.37 (.16)**
PID strength	-1.60 (.27)**	-1.28 (.29)**	-1.20 (.29)**	-1.40 (.34)**	-1.33 (.35)**
Pres. rat.	.26 (.19)	—	—	—	—
House rat.	.04 (.24)	—	—	—	—
Dem. INC.	.16 (.16)	—	—	—	—
PID × Pres. rat.	—	.85 (.21)**	.78 (.22)**	.54 (.26)*	.48 (.26)*
PID × House rat.	—	3.09 (.37)**	3.11 (.37)**	3.06 (.40)**	3.07 (.41)**
PID × Dem. INC.	—	-.67 (.18)**	-.66 (.19)**	-.84 (.22)**	-.80 (.22)**
COGN	—	—	.52 (.19)**	—	.47 (.22)*
DISTANCE	—	—	—	.91 (.30)**	.87 (.31)**
N =	1,062	1,062	1,033	817	795
% CC	78	80	80	82	82
Nagelkerke Pseudo-R ²	.07	.28	.29	.32	.31

Standard Error in parentheses. * $p < .05$; ** $p < .01$.

Notes: Entries are logistic regression coefficients, with standard errors in parentheses. Column 1 is an additive model that does not multiply the last three independent variables by "PID". The dependent variable is a dummy that equals 1 for ticket-splitters and 0 for straight-ticket voters. "PID" is coded 1 for Democratic partisans, -1 for Republican partisans, and 0 for pure independents. "PID strength" is coded 0 for pure independents, .5 for leaners or weak partisans, and 1 for strong partisans. "Presidential rating" and "House rating" are the difference between the number of good points mentioned about Bush (or the House Republican candidate) and the number of good points mentioned about Clinton (or the House Democratic candidate), rescaled 0 to 1. "Dem. INC." is coded 1 if the House incumbent is a Democrat, 0 otherwise. "COGN" is coded 1 for respondents preferring a divided government (or being indifferent) and 0 for those preferring a unified government. "DISTANCE" is coded 0 if ideological distance with the opponent party on the 7-point position scale (where 1 = extremely liberal and 7 = extremely conservative) is small (0-2 points), -5 if distance is medium (3-4 points), and -1 if distance is large (5-6 points).

TABLE 2B
Models of Split-Ticket Voting and Cognitive Madisonian Effects, 1996

	(1)	(2)	(3)	(4)	(5)
Constant	.18 (.28)	-.09 (.26)	-.38 (.30)	-.03 (.27)	-.30 (.31)
Party ID	-.15 (.14)	.19 (.16)	.16 (.16)	.15 (.17)	.14 (.17)
PID strength	-2.16 (.38)**	-1.68 (.38)**	-1.60 (.41)**	-1.39 (.43)**	-1.34 (.43)**
Pres. rat.	-.98 (.27)**	—	—	—	—
House rat.	.42 (.21)*	—	—	—	—
Dem. INC.	-.19 (.25)	—	—	—	—
PID × Pres. rat.	—	.66 (.31)*	.67 (.31)*	.50 (.34)	.53 (.34)
PID × House rat.	—	1.29 (.25)**	1.26 (.25)**	1.46 (.27)**	1.44 (.27)**
PID × Dem. INC.	—	-.63 (.27)**	-.59 (.27)**	-.48 (.29)**	-.46 (.29)
COGN	—	—	.74 (.37)*	—	.67 (.39)*
DISTANCE	—	—	—	.91 (.38)**	.83 (.39)*
N =	610	610	607	557	556
% CC	81	80	80	82	81
Nagelkerke Pseudo R ²	.14	.25	.25	.26	.27

Standard Error in parentheses. * $p < .05$; ** $p < .01$.

Notes: Entries are logistic regression coefficients, with standard errors in parentheses. Column 1 is an additive model that does not multiply the last three independent variables by "PID". See notes to Table 2a for a description of the coding of all variables.

“good points” than Clinton were significantly more likely to split in 1992, as the coefficient for (PID \times Pres) shows (Table 2a, column 2).

The pattern repeats itself in 1996, with the influence of the Dole-Clinton comparisons (Table 2b, column 2). But while presidential candidate evaluation is relevant for partisans, local candidate evaluation is much more so. Look especially at the coefficient of (PID \times House) variable, which interacts party identification with local candidate evaluation. For example, a Democrat who is positive toward a Republican House candidate is much more likely to vote for that candidate, while staying with the Democratic presidential contender. In both elections, this highly significant variable registers the largest partial R (respectively, .25 and .20), indicating that the biggest source of split-ticket voting is defection to the local candidate of the rival party. Incumbency at the local level is also important. For example, if the Democrats field a House candidate who is incumbent, a Republican is more likely to split in his or her favor, as the negative coefficient of (PID \times Democratic Incumbent) indicates. Generally speaking, local evaluations, as they operate in a partisan context, appear key parts in any explanation of ticket-splitting. All told, the model offers a broad structural understanding of ticket-splitting; hence we use it as a backdrop for assessment of cognitive Madisonianism, the central task of this effort.

Demonstrating the Effects of Cognitive Madisonianism

Voters who are cognitive Madisonians want to check and balance political power and governmental policy. If they vote Republican for President, they will be inclined to vote Democrat for Congress, and vice versa. One motivation is to check the concentration of public power in the hands of one party, something viewed as destructive of the common good. Another motivation is to balance policy, on the belief that a mix of Republican and Democratic solutions optimizes the collective benefit (see Fiorina 1988, 430–59.). The first motive has an *institutional goal*: divide political control of the House and the presidency. The second motive has a *policy goal*: provide a blend of Republican and Democratic actions. Since either motive encourages ticket-splitting, it is important to measure them both, and incorporate these measures into the general voting model.

Operationalization of the *institutional goal*—checking power—is straightforward, and employs the single separable preference items already discussed. To enhance interpretation, scores on these indicators are dichotomized (open to divided government = 1, otherwise = 0), then each added as variable COGN to the appropriate baseline model, in column 3, Tables 2a and 2b. In both elections, its coefficient is in the expected direction and statistically significant. Net of partisanship and candidate forces, those who expressed a desire for divided government in 1992 and 1996 appear to have acted on that desire, for they were clearly more likely to split their ticket.

The *policy goal*—balancing party programs—we have neglected thus far, in terms of theory and measurement. We assume that the policy balancers are essen-

tially party moderates, who see the opposition party as ideologically dominated by moderates not so different from themselves. (See also Alesina and Rosenthal 1995 and Mebane 2000 for other variations on the notion of splitters as moderators.) By endorsing the idea of moderates in both parties working together, they believe they can avoid radical extremes and achieve a mix of mainstream Republican and Democratic policies. Split-ticket voting comes about, first, because it is ideologically not much of a leap (that is, they do not see themselves as having that much policy disagreement with their opponents), and, second, because splitting the ballot helps prevent one party takeover, which would allow policies to be pushed to the party extreme, left or right.

Preliminary evidence for this idea comes from comparing the ideological perceptions of splitters versus nonsplitters. Respondents are asked to place themselves, and the Democratic and Republican parties, on a seven-point scale (0–6) from liberal to conservative. Table 3 shows these ideological placements, for splitters and nonsplitters by party identification, in the most recent survey. Observe that splitters, be they Democrat or Republican identifiers, place themselves near the middle of the scale, at almost the same value (4.0–4.5). This stands in sharp contrast to the straight-ticket voters in these parties, who separate themselves by about two points (3.5–5.3). Thus, splitters appear moderate, and more or less equally so regardless of party attachment. Next, observe that splitters see themselves as much closer to the opposition party, as compared to nonsplitters. Republican ticket-splitters see themselves as 1.4 points away from the Democrats (4.5–3.1), whereas Republican nonsplitters place themselves 3.1 points away, over twice that distance (5.3–2.2). The same pattern repeats itself for the Democrats, with splitters only 1.0 away from the Republican perception, but nonsplitters 1.7. Overall, split-ticket voters, irrespective of party stripe, see themselves as ideologically much more like the opposition, than do straight-ticket voters. For these voters, who are moderates, it might be relatively easy to cross party lines, in search of a moderate coalition against partisan extremes.

In general, then, such balancing behavior appears more likely when a partisan's own ideological position is not far from his or her perception of the oppo-

TABLE 3

Ideological Self-Position and Perceptions of Parties Among
Ticket-Splitters and Straight-Ticket Voters, 1996

	Perception of the Democratic Party	Self-position	Perception of the Republican Party
Democrats (straight ticket)	3.3	3.5	5.2
Democrats (ticket splitting)	3.3	4.0	5.0
Republicans (ticket splitting)	3.1	4.5	5.0
Republicans (straight ticket)	2.2	5.3	5.4

Note: Entries are mean scores on a seven-point ideological position scale, where 0 = extremely liberal and 6 = extremely conservative.

sition party's ideological position. Operationalization of the policy goal of balancing, so conceived, leads to an uncomplicated use of the placement scores on the seven-point liberal-conservative scale. Imagine Democrat A gave a self-placement of "2" and placed the Republican party at "6," yielding an ideological distance score of $|2 - 6| = 4$. Democrat B, in contrast, gave a self-placement of "4" and placed the Republican party at "5," yielding an ideological distance score of $|4 - 5| = 1$. This variable, labeled DISTANCE, is composed of absolute values from "0" to "6." Those with lower distance scores (e.g., Democrat B compared to Democrat A) are closer to the rival party and thus more likely to seek policy balance, voting for the opposition party in one of the two races. Bivariate analysis shows a clear relationship between distance and split-ticket voting.

Does the suggested relationship survive multivariate control? Yes, according to column 4, Tables 2a and 2b, where the above distance variable is added to compete the model specification. Interestingly, the coefficient is almost the same magnitude in 1992 and 1996 and is easily significant (at about the same level) in both. Partisans who see the opposition as little different from themselves on the issues seem more likely to split their vote, presumably to avoid ideological extremes and achieve a more moderate policy balance.

Our two measures of cognitive Madisonianism, one the checking of power, the other the balancing of policy, stand on solid empirical ground. In both 1992 and 1996, the effects of these variables are statistically significant, even after controlling on strong party and candidate influences. (Further, the statistical significance is sharper still, if the two surveys are pooled into one). The results clearly support the argument for intentional ticket-splitting. Voters who are open to divided government, and who see themselves as not very ideologically distinct from the opposition party, are more likely to split the ticket. These cognitive Madisonians seem motivated by the institutional goal of checking power, and the policy goal of balancing policy.

How Big Are the Effects? A Look at the "Model Madisonian"

It seems established that cognitive Madisonianism was a factor in the ticket-splitting of 1992 and 1996. But how big a factor? Simply showing statistical significance does not directly address its substantive significance. How likely are cognitive Madisonians to split their ticket, compared to other voters? As a baseline, we know that the probability of a split from a voter drawn at random in the 1992 survey is .22 (the sample mean, the proportion who said they did vote a split-ticket). The comparable figure for 1996 is .20. A convenient way to proceed is to calculate the changes in the probability from a change in variables evaluated at these sample mean values (see Petersen 1985, 131). We thus vary the cognitive and distance variable values from one extreme to the other, compute the predicted log odds of a split, exponentiate and transform to vote probabilities (see DeMaris 1992, and Menard 1995, 12–13).

If a voter favors unified government and is not ideologically close to the opposition party, then the probability of a ticket-split is negligible, at only .07 for 1992

and .05 for 1996. But suppose a voter is a “Model Madisonian.” That is, he or she is open to divided government (score = 1 on COGN) and is ideologically close to the opposition party (score = 0–2 on DISTANCE). Under those conditions, the probability of a ticket-split jumps to .52 in 1992, and .56 in 1996. Clearly, the presence of Madisonian thinking can heavily influence the ticket-splitting behavior of the individual voter. Moreover, these individual level effects are capable of generating substantial aggregate shifts, especially when the percentage of “Model Madisonians” in the electorate is not small. In the 1992 survey, 26% of the respondents who said they voted (split ticket or not) could be so labeled. (Further, this estimate seems free of bias from balancers of “conditional sincerity,” who voted straight-ticket to achieve division). In the 1996 survey, the comparable figure is 20%. Last, but not least, these results reveal the micro-foundations of the aggregate shift from the 1992 Republican President/Democratic Congress pattern to the 1996 Democratic President/Republican Congress pattern.

Comparison to Rival Perspectives

Our evidence suggests that ticket-splitting can be intentional, and thus goes against two major bodies of work, that inspired by Jacobson (1990) and that inspired by Sigelman et al. (1997). Let us consider the earlier tradition first. Jacobson argues that “Presidential candidates are evaluated according to their views on national issues and their competence in dealing with national problems . . . Congressional candidates are evaluated on their personal character and their devotion to district services and local issues” (1990, 115). The electorate in its majority tends to think Republicans can handle national issues better, while Democrats can handle local issues better. Thus, divided government occurs as an unintended side-effect of voters exercising these fundamental preferences in the two contests.

This insight into the electoral origins of divided government is profound, and we cannot do it justice in our brief article. However, we do believe that the essence of the Jacobson argument is captured, and thus controlled for, in the full model specification, column 5 (Tables 2a and 2b). As we understand the Jacobson argument, what counts for voters are candidate stands on issues, particularly in the context of their party affiliation; in other words, what is measured with our presidential candidate and local candidate evaluation variables. With each measure, the voter compares the Republican and Democratic candidates in terms of their “good points.” For example, suppose a Democrat believed that Dole was better on national issues than Clinton, and thus gave Dole more “good points,” split-voting for him in 1996. That scenario, which could also be applied to the House, seems to fit with the Jacobson idea, and is specified in the baseline model. Hence, our divided government results reveal the part of the voter decision-making processes that lie beyond the underlying majority Republican and Democratic institutional preferences at work.

Let us now consider the null findings on divided government preference and split-ticket voting, represented in the Sigelman et al. (1997) analysis of the 1992 ANES. Why do their results flatly contradict our positive ones? The essential reason is because they employ multinomial logit in a way which makes interpretation highly problematic. In particular, their null findings are a product of the peculiar reference category they select for the logit. It is perhaps worth explication of this methodological point. Here is the set of equations they estimate, with the probability of split Republican president to Democrat House (rd) or split Democrat president to Republican House (dr) noted in brackets along with the probability of no split (rr, dd):

$$\log[P(\text{rd})/P(\text{rr})] = a_1 + Xb_1 \quad (1)$$

$$\log[P(\text{dr})/P(\text{rr})] = a_2 + Xb_2 \quad (2)$$

$$\log[P(\text{dd})/P(\text{rr})] = a_3 + Xb_3 \quad (3)$$

As can be seen, the dependent variable is considered to have $G = 4$ categories (rd, dr, rr, dd), so requiring $G - 1 = 3$ equations. While only three independent equations can be constructed, altogether $G(G - 1)/2 = 6$ different odds can be predicted (DeMaris 1992, 62). The choice of odds for examination is extremely important. More to the point, “As the choice of category for the denominator of these odds is arbitrary, it becomes incumbent upon the analyst to choose a category that facilitates interpretation” (DeMaris 1992, 65). Note that Sigelman et al. (1997, 884) chose (rr) for the denominator, so posing great difficulties in interpretation.

Consider the meaning of each logit in the above equations. Begin with equation 3, where the difficulty is especially clear. Equation 3 predicts how the log odds of straight-ticket voting, (dd) versus (rr), will change in response to changes in the independent variables. In other words, it seeks to explain why some voters always vote Republican, while others always vote Democrat. It is not surprising that cognitive Madisonianism would have no effect on that choice. Equation 2 asks about the probability shift in voting straight Republican (rr) versus Republican for the House only (dr). That is a rather unusual contrast, essentially comparing strong Republican partisans (who largely compose the rr category) to the modal split group (dr) for Democratic partisans (recall Table 2). Again, it is not surprising that cognitive Madisonian variables fail to distinguish between these groups. Finally, equation 3 contrasts straight-ticket Republican voters (rr) with Democratic defections to the House (rd). This is a more plausible comparison, given we know this is the modal split for Republican identifiers. (Interestingly, it is for this contrast that divided government attitudes come closest to having statistically significant effects; see Sigelman et al. 1997, 888). However, respondents in these two categories—(rr) and (rd)—compose only 42% of the sample, so the cognitive Madisonian effects are swamped by noise. For multinomial logits, the prediction is that the respondent falls in one of the two constructed categories

(Theil 1970). We argue that the preferred constructed reference category, for the logit denominator, is simply straight-ticket voters (rr, dd), which is what we used.

We are not the first to find divided government effects, as our review of current literature, especially that on nonseparable preferences, makes clear. However, we are perhaps unique in finding rather strong effects. One reason for that is inspiration from an essential idea of the nonseparable preferences school. Smith et al. (1999), and others, argue that it is not enough to say that Democrats vote for a Democratic president, because that vote depends on the partisan character of the House race. Put another way, preferred partisan outcome in one dimension is conditional on partisan calculations in another dimension. Our argument is not this, but it has an important similarity: preferred partisan outcome in one dimension is induced by candidate evaluation, conditional on the social-psychological dimension of party identification. That specification change, leading to an additive-nonadditive formulation of the process, much improved our model. Thus, we agree with the path-breaking methodological point made by Smith et al. that “it is not enough to posit unconditional, noninstitutional measures as the sole determinants of partisan preference” (1999, 756).

Besides offering a new model specification, we also measured the policy balancing idea differently from past work. The distance measure is operationally simple, (e.g., a Democrat’s own ideology score minus the perceived Republican score). Further, it is conceptually simple, and that is in keeping with what appear to be the information processing limits of most voters (Delli Carpini and Keeter, 1996). The notion seems within the grasp of the average citizen—partisans may defect when they see the opposition as not much different on the issues.

Summary and Conclusions

American voters who favor divided government *do* appear to act on that preference, by voting a split-ticket. In 1992 and in 1996, voters who declared themselves open to different party control across electoral arenas were clearly more likely to vote one party for the presidency, another for Congress. That is, they voted to check governmental power, by dividing it between two institutions. Voters who say they want to split control tend to vote that way, and that relationship holds up under extensive testing. American voters who were policy balancers, especially those moderates who see themselves as ideologically not for the opposition, also tended to split the ticket. By voting for the opposition party, they sought to curb party extremes and obtain a moderated policy mix.

Overall, the case for the effects of the two components of cognitive Madisonianism—checking power and balancing policy—seems strong. In two very different elections, there existed a body of voters who weighed, in some way, the pros and cons of party division, contemplating its impact on power distribution and policy outputs. They went on to vote a split ticket, in order to achieve their goals. Among “Model Madisonians,” those who value divided government and see little ideological difference between themselves and the opposition party, the

probability of splitting the ticket is quite high. Moreover, this model group was not small, numbering around a quarter of the electorate in both 1992 and 1996. Hence, the microfoundations of divided government appear pervasive and strong in the electorate. Tracing their variations should go a long way toward explaining changing patterns at the aggregate level of national institutions.

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