DO LEGISLATORS PAY TO DEVIATE FROM CONSTITUENTS?

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INTRODUCTION

A large literature analyzing roll call votes has studied legislators' deviations from constituents' preferred policies. Beginning with Kau and Rubin [1979], the standard empirical technique has been to regress a legislator's vote record on his own characteristics and variables that measure his district's economic interests. The residuals from this type of regression are commonly viewed as a measure of legislators' deviations from their constituents' predicted policy preferences. Frequently these residuals are significant predictors of subsequent roll call votes. To interpret this explanatory power, the literature has variously identified these residuals as legislator shirking, legislator ideology, legislator signaling as in an agency model, and so on.

We examine the empirical relationship between vote score residuals and incumbents' campaign expenditures in the subsequent re-election bid. The incumbent's goal is to get re-elected. The literature on legislator voting has consistently shown that deviating from constituents' preferred policies makes this objective more difficult [Bender and Lott, 1996]. Despite this, an incumbent may occasionally deviate from his constituents' preferences with good reason (to pursue his personal views, to service interest group pressure, to demonstrate loyalty to party leaders, etc.). To deviate while diminishing the costs of deviating, incumbents may increase their campaign expenditures in an effort to convince constituents to vote for them regardless of their policy positions. This interpretation of re-election expenditures has been termed *persuasive* campaigning [Mueller and Stratmann, 1994]. If we view congressional seats as an economic good and re-election expenditures as the unit price paid by an incumbent for the good, incumbents may "buy" some deviation, in the form of diminished electoral punishment, by expending more in their re-election bids. Relying on this reasoning, we expect incumbents' deviations to be positively related to their subsequent campaign expenditures. Our results reveal a statistically signifi-

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cant positive empirical relationship, other things equal, for the 1992 through 1998 Congressional elections.

In the next section we discuss the theoretical motivation of the paper and continue by presenting our empirical model in detail. We then discuss the results, and finish by offering our concluding remarks.

THEORETICAL CONSIDERATIONS

Consider a one-dimensional policy space on which the median voter in the k^{th} congressional district in year t establishes position M_{kt}. The legislator representing the k^{th} district, through his vote record in year t, establishes position L_{kt}. Hence, the legislator's deviation from the median constituent can be written as the absolute difference between the positions $\mu_{kt} = |L_{kt} - M_{kt}|$. Unlike previous studies, we are interested only in the *degree* of deviation rather than the direction.

After Kau and Rubin [1979] initially interpreted the real value of μ_{kt} as a measure of the legislator's own ideological views, a large literature developed to explain the nature of these residuals. Other early contributors interpreted it as a measure of "shirking" or of "ideological consumption" in a model where legislators are imperfectly monitored agents of voter-principals [Kalt and Zupan, 1984]. It was argued that this undermined a rational choice approach to politics, even though the "consumption" of μ_{kt} appeared to decrease with proxies of the cost of shirking [Nelson and Silberberg, 1987, Kalt and Zupan, 1990]. Another interpretation is that deviations from constituents are signals of the legislator's ideological convictions—a way for constituents to economize on the costs of assessing and monitoring the productivity of alternative legislators. This view considers μ_{kt} as an investment in ideological reputation [Lott, 1987; Dougan and Munger, 1989]. Still other possibilities lie outside the signaling and shirking alternatives. Legislators may deviate from constituents for purposes of servicing interest groups and acquiring campaign contributions, or to enter logrolling agreements, or to satisfy party leaders for the purposes of acquiring better committee seats [Leighton and López, 2002]. Given the many possibilities in this literature, interpreting μ_{kt} seems beyond the capabilities of a single study. Our approach avoids this delicate task.

Instead, we focus on the idea that deviating from the median voter's ideal point, toward the legislator's own ideal point, is presumably of some value to the legislator—regardless of how μ_{kt} is conceived. We are most interested in whether the incumbent perceives such deviations as making his re-election efforts more difficult. The literature on congressional voting is nearly unanimous in showing that when legislators do deviate from constituents, they are punished fairly quickly at the voting booth.[1] Considering these findings, we ask whether incumbents who deviate more tend to "buy back" the resultant electoral losses by spending more in their re-election bid.

Consider the problem in a simple price-theoretic model. If a legislator values movement away from M_{kt} toward L_{kt}, then $\mu_{kt} = |L_{kt} - M_{kt}|$ would be one argument in a legislator's demand function for his congressional seat. If we hold other arguments constant, we may define $q^k[p(\mu_{kt})]$ as the k^{th} legislator's non-negative demand for a congressional seat. The argument (p) indicates the unit price paid to win re-election—that is, the incumbent's re-election campaign expenditures. The "price" is a function of the degree of deviation. If this price-theoretic view is an accurate representation of the incumbent's re-election calculus, we expect greater deviation to increase the unit price paid for the seat, other things equal. That is, $\partial p/\partial\mu > 0$.

The data that we use are directly amenable to this theoretical construction. We use the legislator's vote record, as captured by ADA scores (tracked by the public-interest group Americans for Democratic Action), to establish L_{kt}. ADA scores are calculated along a single left-right (liberal-conservative) dimension, as in our model. We use Census data on constituents' economic and demographic characteristics to establish M_{kt}. While these data are on multiple dimensions, we regress L_{kt} on a set of variables measuring M_{kt}. We then take the residuals, which are along the same liberal-conservative dimension as our measure of legislator deviation. Hence, our data fit the single dimension of this model conveniently.

Our approach also introduces a useful further implication. We use the incumbent's campaign expenditures to measure (p), and the viability of this concept depends on what function campaign expenditures serve. Are campaign expenditures primarily informative, in that they inform voters of the legislator's position? Or are campaign expenditures primarily persuasive, reflecting the incumbent's attempt to persuade voters to vote for him regardless of the deviation? Mueller and Stratmann [1994] demonstrate within a one-dimensional theoretical model of elections that campaign expenditures are almost always persuasive. While we do not present a model of re-election because our empirical approach avoids estimating whether the incumbent or challenger wins, our results can shed empirical light on the theoretical model in Mueller and Stratmann [1994]. Our hypothesis that incumbents can reduce electoral punishment for their deviations by increasing campaign expenditures implies that incumbents' expenditures are primarily persuasive rather than informative. Therefore, evidence for our hypothesis also supports the hypothesis of persuasive campaign expenditures.

EMPIRICAL SPECIFICATION

Our measure of deviation from constituent interests is provided by the vote score residualization technique of Kau and Rubin [1979] and Kalt and Zupan [1984; 1990]. These authors were originally concerned with developing a proxy for legislator ideology and estimating the explanatory power of that proxy in a model of legislator voting. They wanted to demonstrate whether a strict rational choice model fit voting data, or whether a model that includes ideology should be used. These contributions were followed with an empirical literature that developed along multiple margins: whether residuals measure ideology or shirking or both, whether legislators shirk, the conditions under which they will shirk more or less, whether and to what extent constituents punish shirking, and so on. Bender and Lott [1996] provide a useful survey of this large literature. For present purposes, we note that many of these contributions use some variation of Kau and Rubin's [1979] measure of deviation from constituent interests: the extent to which the legislator's vote record is unexplained by measures of constituent economic interests. To illustrate the technique, and our variation on it, let **Y** represent a matrix of i roll call votes by k legislators, which would be estimated using some variation of the following general model:

$$\mathbf{Y} = \alpha_1\mathbf{D} + \beta_1\mathbf{V} + \eta \tag{1}$$

Here, **D** defines the relevant economic interests of the k congressional districts, while **V** is an index of legislators' vote records *prior* to voting on **Y**, the issue at hand. ADA scores are the most commonly used index. Kau and Rubin argued that these vote records contain information about legislator ideology. If the elements of β_1 are statistically significant, then ideology may influence legislator voting. But **V** and **Y** represent the same underlying functions—the legislator's voting calculus. Hence, this regression would produce biased estimates of both α_1 and β_1. To solve this problem, Kau and Rubin introduce a two-step regression technique—first regressing economic interests and exogenous variables on past voting, and then using the *unexplained* portion of the past voting (the "non-economic" portion) as a measure of ideology to be regressed on current voting. That is,

$$\mathbf{V} = \pi_1\mathbf{D} + \pi_2\mathbf{C} + \mu \tag{2a}$$

$$\mathbf{Y} = \alpha_2\mathbf{D} + \beta_2\hat{\mathbf{u}} + \eta. \tag{2b}$$

In equation (2a), **C** is a matrix of exogenous constituent characteristics that are not necessarily affected by the outcome on **Y**. The residuals from this equation, $\hat{\mathbf{u}}$, now represent the portion of the vote record that goes unexplained by legislator and district interest variables. Kau and Rubin [1979] interpret these residuals as a measure of legislator ideology. In regressions like equation (2b), the residuals $\hat{\mathbf{u}}$ are statistically significant and robust predictors of legislator voting.[2]

The literature that followed nearly exclusively uses roll call votes as the dependent variable in equation (2b). Our interest, however, lies not in predicting an incumbent's vote on a particular issue, but in whether deviations from constituent interests, measured by the residuals $\hat{\mathbf{u}}$, motivates increased incumbent expenditure in an effort to diminish electoral punishment. Accordingly, we depart from convention in the empirical voting literature and use incumbent re-election expenditures as our dependent variable instead. We estimate:

$$\mathbf{V} = \pi_1\mathbf{D} + \pi_2\mathbf{C} + \mu \tag{3a}$$

$$\mathbf{E} = \alpha_3\mathbf{X} + \beta_3\hat{\mathbf{u}} + \varepsilon. \tag{3b}$$

Equation (3a) is the same as equation (2a); we generate the residuals in the same fashion as the literature. In equation (3b), however, we estimate incumbent campaign expenditures, **E**, using a standard empirical campaign spending model, **X**, while adding the residuals $\hat{\mathbf{u}}$ to investigate our main empirical inquiry. Hence, we generate the residuals in the conventional way, but then use them in a model of campaign spending rather than a model of legislator voting.

We are aware that equation (2) has been the subject of much criticism in the empirical legislator voting literature. The econometric critiques include underspecification of the **D** matrix [Peltzman, 1984; 1985], omitted variables bias on β_2 estimates, and endogeneity bias on α_2 estimates [Jackson and Kingdon, 1992]. It is reasonably clear that equation (2b), in some sense, attempts to explain votes with votes. Furthermore, to attach a "signaling" or "ideological shirking" interpretation to μ, an unobservable, seems to require an intellectual leap.[3] Our measurement of **D** follows standard procedure of including average economic and demographic variables by congressional district. This is limited by the data, and the number of variables we can include is limited to avoid the omitted variables critique. The more serious concern is with bias due to explaining votes with votes. We avoid this by modeling expenditures in equation (3b) instead of current votes in equation (2b). Finally, we do not specifically interpret the residuals $\hat{\mathbf{u}}$; instead we treat them simply as what they are—deviations from perceived constituent interests. By removing ourselves from equation (2), avoiding the interpretation of $\hat{\mathbf{u}}$, and instead modeling campaign spending as in equation (3), we distance ourselves from the above criticisms. Again, while we generate the residuals in the conventional manner, we use them in an innovative way.

EMPIRICAL APPLICATION

Our data set contains vote records, campaign information, and legislator characteristics for all 1,511 House incumbents seeking re-election in 1992, 1994, 1996, and 1998.[4] We also observe constituent interest variables for each of these incumbents' districts in each year. The vote record is provided by the annual ADA score, which measures the frequency (on a 0-100 scale) of a legislator voting consistently with the Americans for Democratic Action's position on a set of key votes during the year. The Americans for Democratic Action selects the votes each year and takes a position on each vote that is easily identifiable as liberal, so a higher ADA score indicates a more liberal voting record. Since we are making intertemporal comparisons of ADA scores, while the set of votes used to assign ADA scores is not constant over time, we have adjusted the raw ADA scores using the linear transformation method from Groseclose, Levitt, and Snyder [1999]. If the k^{th} legislator's raw ADA score in year t is y_{kt} then the adjusted ADA score is $\hat{y}_{kt} = (y_{kt} - a_t)/b_t$ where a_t and b_t are maximum likelihood parameter estimates from Groseclose, Levitt, and Snyder [1999]. The a_t and b_t estimates provide an index for converting raw ADA scores in any year to adjusted ADA scores for that year—similar to how a price index is used to convert nominal to real economic variables.[5] The campaign spending data are from Federal Election Commission databases [FEC, 2002],[6] while the legislator-specific variables come from the *Almanac of American Politics* [Barone and Ujifusa, 1994–2000]. Constituent and/or district variables come from 1990 Census data. Table 1 lists all variables used in this study, with definitions, sources, and descriptive statistics. We would prefer to use panel estimation methods, but many of our independent variables are taken from 1990 Census data and do not vary over our sample years. Therefore, we treat the data as a pool of four cross-sections over time and use both year and state dummies to capture possible intercept effects.

TABLE 1
Variable Names, Definitions, Sources, and Descriptive Statistics

Variable Name	Description (Source)	Units	Mean	Std. Dev.	Min.	Max.
EXPEND	Incumbent Campaign Expenditures[a]	$1000s	582.8	436.9	6.62	6541.3
ADJADA	Incumbent's ADA score (Groseclose et al. 1999)	None	45.9	36.4	−7.5	102.0
ABSRESID	Absolute value of residuals (Table 2)	None	.848	.529	.0005	2.71
CHSPEND	Challenger Campaign Expenditures[a]	$1,000s	184.7	301.1	0	3325.9
CHSPEND-IV	Instrument for CHSPEND (Table 3)	$1,000s	228.5	169.4	−383.5	937.0
VOTEMARG	Incumbent less challenger vote share[a]	Percent	33.1	25.5	−100	100
VOTEMARG-IV	Instrument for VOTEMARG (Table 3)	$1,000s	33.1	15.2	−36.3	69.4
TENURE	Incumbent's years in office[b]	Years	9.54	7.74	1	51
GENDER	=1 if incumbent is female[b]	{0,1}	.104	.305	0	1
AGE	Age of Incumbent[b]	Years	53.2	9.70	28	87
MAJPARTY	=1 if incumbent is in majority party[b]	{0,1}	.572	.498	0	1
HHIE	Economic Concentration (Authors' calculations)[d]	{0−10,000}	200	48.9	76.69	381.3
PCTURBAN	Population in Urban Areas[c]	Percent	63.9	31.5	0	100
PCTFARM	Population on Farms[c]	Percent	1.53	2.27	0	14.2
PCTPUBEMP	Total Public Employees[c]	Percent	2.21	1.20	.63	11.2
PCTBLACK	Race of Constituents[c]	Percent	11.8	15.9	0	73.9
PCTCOLLG	Education of Constituents[c]	Percent	18.3	6.56	5.3	48.3
MEDINCM	Income of Constituents[c]	$1,000s	30.858	8.326	15.052	57.219
WELFARE	Welfare spending in District[c]	$1,000s	65415	40798	13572	308223
CORPCONT	Corporate Contributions to Incumbent[a]	$1,000s	99.602	84.797	−3.000	624.547
LBRCONT	Labor Contributions to Incumbent[a]	$1,000s	55.197	68.694	−3.000	408.217
STATEFUND	= 1 if state finances elections (Council of State Governments)	{0,1}	.536	.498	0	1
CHBEGIN	Challenger beginning cash[a]	$1,000s	3.54	29.56	0	670.9
CHCORP	Challenger corporate contributions[a]	$1,000s	7.74	25.95	−.5	396.2
CHLABOR	Challenger labor contributions[a]	$1,000s	15.73	39.69	−1	2.32
OPENPRIMARY	= 1 if open or blanket primary election (Westley et al. 2002)	{0, 1}	.515	.499	0	1
PEROT92	Percent of District voting Perot in 1992[b]	Percent	18.32	5.99	3	33
WINNER	= 1 if Incumbent Wins	{0, 1}	.937	.241	0	1

a. Federal Election Commission (various years)
b. Barone and Ujifusa (various years)
c. United States Census Bureau (various years)
d. HHIE is a Herfindahl-Hirschman index on U.S. Census data for employment in seventeen industries. The higher the index, the more concentrated is the employment in certain industries, indicating more concentrated economic interests and greater political effectiveness in the district, ceteris paribus. HHIE varies by district from a low of 77 (New York, 16th) to a high of 381 (California, 15th). The former occupies most of the Bronx, and is the most economically diverse district in the country according to this index. The southeastern part of Kentucky is similarly diverse, scoring a mere 80 HHIE. The latter is California's Silicon Valley around the city of San José, a presumably economically concentrated region. Similarly, Virginia's 8th District around the Pentagon in Arlington scores a 363, and New York's 14th District, the east side of Manhattan, scores a 360.

TABLE 2
OLS to Obtain ADA Residuals
Dependent Variable is ADJADA

	Coefficient Estimate	Standard Error
AGE	−0.167[c]	.098
TENURE	0.563[a]	.125
MAJPARTY	−3.64[b]	1.55
GENDER	15.33[a]	2.65
PEROT92	−0.162	.180
PCTURBAN	27.62[a]	4.07
PCTFARM	36.07	48.6
PCTBLACK	27.13[a]	7.51
PCTCOLLG	13.40	20.4
PCTPUBEMP	193.69[b]	69.2
HHIE	0.173[a]	.025
Ln(*MEDINCM*)	−27.89[a]	5.86
Ln(*WELFARE*)	26.22[a]	1.88
1994−Dummy	−1.90	2.22
1996−Dummy	−8.79[a]	2.23
1998−Dummy	−6.79[b]	2.21
Constant	−5.06	62.94
F(16, 1493) = 47.11		
Prob > F = 0.00		
$\bar{R}^2 = .328$		

Dependent variable is ADA scores that have been adjusted for intertemporal comparison as explained in text.
Sample is House incumbents seeking re-election in 1992, 1994, 1996, and 1998 (N = 1,511).
See Table 1 for definitions, sources, and descriptive statistics of all variables.
a. Significant at the 99 percent confidence level.
b. Significant at the 95 percent confidence level.
c. Significant at the 90 percent confidence level.

Our first step is to obtain the vote score residuals. As in equation (3a) above, let the adjusted ADA score for the k^{th} legislator in year t be denoted as V_{kt}. The residuals are obtained by estimating (3a), and μ_{kt} represents the vote score residual. Table 2 presents the results from this estimation.

The variables in **D** and **C** from (3a) account for the explained portion of the past vote record. Since these variables consist of the potential economic interests affected by the vote record, the unexplained portion, μ_{kt}, is called the "non-constituent" portion of the vote record. A positive value for μ_{kt} indicates that the k^{th} legislator voted more liberal than the district's median voter in year t, and a negative value indicates a deviation in the conservative direction. As discussed earlier, we are interested only in the *degree* of deviation, not the direction. Therefore, we convert these residuals to their absolute values and assign the variable name *ABSRESID*, which is our explanatory variable of primary interest in the next estimation.[7] The summary statistics for *ABSRESID* appear in Table 1.

Our next step is to estimate equation (3b), the incumbent expenditures equation. A well-known simultaneity problem exists in models of campaign spending. With incumbent spending as the dependent variable, both challenger spending and margin of victory are theoretically significant variables. It is also expected, however, that incumbent spending, challenger spending and the vote margin are endogenously determined. Intuitively speaking, as a race becomes tighter and the challenger spends more, the incumbent will be motivated to spend more as well. At the same time, as the incumbent spends more, he or she will increase the margin of victory, which may in turn impel the incumbent to spend less. A Hausman test using the current data indicates that challenger spending is unambiguously endogenous, while victory margin is less clearly endogenous.[8] We approach this problem using the following three-equation system:

(4a) $$E_{kt} = \beta_0 + \beta_1 ABSRESID_{kt} + \beta_2 \hat{CHS}_{kt} + \beta_3 \hat{MAR}_{kt} + \beta_j X_{jkt} + \varepsilon_{kt},$$

(4b) $$\hat{CHS}_{kt} = \gamma_0 + \gamma_1 ABSRESID_{kt} + \gamma_j X_{jkt} + \theta_C \mathbf{Z}_{Ckt} + r_{Ckt}, \text{ and}$$

(4c) $$\hat{MAR}_{kt} = \delta_0 + \delta_1 ABSRESID_{kt} + \delta_j X_{jkt} + \theta_M \mathbf{Z}_{Mkt} + r_{Mkt}.$$

The subscripts k and t indicate congressional district and year, as before, and j indicates the j^{th} exogenous explanatory variable. Equations (4b) and (4c) are the reduced-form expressions for the endogenous explanatory variables, namely challenger spending and margin of victory, respectively. Hence, Z_C and Z_M are the instrumental variables (that is, exclusion restrictions) used in the first stage equations. Following standard practice, we assume $E(r_C) = E(r_M) = 0$ and $Cov(r_C, \varepsilon) = Cov(r_M, \varepsilon) = 0$. We do not assume independence between r_C and r_M. As with all two-stage least squares (2SLS) estimations, the quality of $\mathbf{Z}_C$ and $\mathbf{Z}_M$, which are subject to data limitations, determine the value of this correction procedure over ordinary least squares (OLS) estimates. We first describe the variables, then discuss the equations—including quality of the instruments—and then proceed to the second stage results.

The endogenous variables are:

- *EXPEND*: total expenditures by the incumbent;
- *CHSPEND*: total expenditures by the incumbent's foremost challenger; and
- *VOTEMARG*: incumbent's vote share less the challenger's vote share (this takes a negative value when the challenger wins).

Our instrumental variables for *CHSPEND* are:

- *STATEFUND*: a binary variable coded 1 for states that provide government funding of statewide elections;
- *CHBEGIN*: challenger's cash on hand at beginning of election campaign;
- *CHCORP*: total corporate contributions to the challenger; and
- *CHLABOR*: total labor union contributions to the challenger.

With *STATEFUND* we are attempting to capture cross-state variation in challengers' campaign finance. Public financing of elections typically includes qualifying thresholds and similar criteria. Since these are typically easy for incumbents to reach but less so for challengers, this may be a variable that explains *CHSPEND* but is not strongly correlated with *EXPEND*. The other variables represent the fundraising successes of the challenger.

Our instrumental variables for *VOTEMARG* are measured as:

- *OPENPRIMARY*: a binary variable coded 1 for states that allow open or blanket primary elections;
- *PEROT92*: percent of a district's vote going to H. Ross Perot in the 1992 presidential election; and
- *WINNER*: binary variable coded 1 for an incumbent who is reelected.

With these variables we attempt to measure effects on the margin of victory distinct from their having an effect on *EXPEND*. States with open/blanket primaries typically have more highly contested races. *WINNER* adds explanatory power, and we use Perot's vote share to proxy coattail effects.

Finally, in the exogenous $\mathbf{X}_j$ matrix we include a typical campaign spending model with variables measured as:

- *MAJPARTY*: binary variable coded 1 if the legislator is a member of the majority party. For example, for Republicans it equals 0 in 1992 and 1994, but 1 in 1996 and 1998;
- *TENURE*: number of years the incumbent has been in the House;
- *GENDER*: binary variable coded 1 if incumbent is female;
- *HHIE*: a measure of economic concentration by district. See Table 1 for full explanation;
- *PCTBLACK*: district race variable, as listed by the Census Bureau;
- *PCTCOLLG*: percent of the district population with a college degree; and
- ln(*MEDINC*): natural log of the district's median income.

The first-stage results appear in Table 3. On *CHSPEND* the instrumental variables are jointly significant and all but *STATEFUND* is individually significant. The adjusted R^2 is a healthy 0.314 and the correlation coefficient between *CHSPEND* and *CHSPEND–IV* is 0.566. In the estimation of *VOTEMARG* the instrumental variables are jointly and individually significant. The adjusted R^2 is 0.314 and the correlation coefficient between *VOTEMARG* and *VOTEMARG–IV* is 0.565. From these results, it appears that the quality of these instruments is not a serious concern.

Moving to the second-stage estimations, we use White's robust standard errors, and, because we have multiple years for each congressional district in the same pooled cross section (typically with the same person as the representative), we also cluster the observations by district to control for variance dependence within districts. In a pool of cross-sections over time, there will likely be intercept effects to be

TABLE 3
First Stage Estimations of Instrumental Variables
Dependent Variable is CHSPEND and VOTEMARG

	Estimation of CHSPEND Coefficient Estimate	Estimation of CHSPEND Standard Error	Estimation of VOTEMARG Coefficient Estimate	Estimation of VOTEMARG Standard Error
ABSRESID	23.76[c]	12.89	−4.28[a]	1.07
MAJPARTY	12.95	13.13	−8.09[a]	1.10
TENURE	−1.59[c]	.862	.195[b]	.070
GENDER	79.73[a]	21.72	.105	1.81
HHIE	.167	.202	−.039[b]	.017
PCTBLACK	−58.23	46.59	15.93[a]	4.96
PCTCOLLG	180.4	162.7	31.77[b]	13.84
Ln(*MEDINCM*)	53.9	43.92	−3.98	3.71
PCTPUBEMP	614.8	570.7	−66.19	47.7
STATEFUND	−.022	13.23		
CHBEGIN	−.916[a]	.237		
CHCORP	4.12[a]	.271		
CHLABOR	3.04[a]	.167		
WINNER			41.43[a]	2.26
PEROT92			−.731[a]	.125
OPENPRIM			−3.29[a]	1.11
Constant	−541.5	421.0	58.43[c]	35.41
	F(13, 1497)	= 54.23	F(13, 1498)	= 58.59
	Prob > F	= 0.00	Prob > F	= 0.00
	$\bar{R}^2$	= .314	$\bar{R}^2$	= .314

Sample is House incumbents seeking re-election in 1992, 1994, 1996, and 1998 (N=1511).
See Table 1 for definitions, sources, and descriptive statistics of all variables.
a. Significant at the 99 percent confidence level.
b. Significant at the 95 percent confidence level.
c. Significant at the 90 percent confidence level.

captured by year dummies. We also check for effects using state dummies. Our rationale in using state (not district) dummies is twofold. First, any variation across districts in addition to that which is explained by **X** and **û** is already controlled for by the clustering of observations (note that this leaves 435 degrees of freedom rather than 1151). Second, because state delegations work together in various capacities, and many campaign and elections laws are determined at the state level, we expect more of an effect across states than congressional districts. To illustrate, we report our findings with and without the state dummies. Finally, because our test for endogeneity is somewhat ambiguous on *VOTEMARG*, we also report results for each of the following cases: (a) treating both *CHSPEND* and *VOTEMARG* as endogenous; (b) treating only *CHSPEND* as endogenous; and (c) treating neither as endogenous—that is, OLS.

Our main findings appear in Table 4. Here we can see that, for all estimations, more spending by the challenger and a narrower margin of victory are both associated with greater spending by the incumbent, other things equal. Also, being in the

TABLE 4
Second Stage Estimations
Dependent Variable is Incumbent's Campaign Expenditures (EXPEND)

	2SLS with Two Endogenous RHS Variables: Model 1	2SLS with Two Endogenous RHS Variables: Model 2	2SLS with One Endogenous RHS Variable: Model 3	2SLS with One Endogenous RHS Variable: Model 4	OLS: Model 5	OLS: Model 6
CHSPEND	—	—	—	—	.300[a]	.285[a]
					(.064)	(.078)
CHSPEND-IV	.357[a]	.343[a]	.308[a]	.304[a]	—	—
	(.074)	(.075)	(.065)	(.069)		
VOTEMARG	—	—	−4.47[a]	−4.51[a]	−3.13[a]	−3.23[a]
			(.514)	(.565)	(.652)	(.711)
VOTEMARG-IV	−5.81[a]	−5.91[a]	—	—	—	—
	(1.09)	(1.19)				
ABSRESID	51.21[a]	57.6[a]	59.11[a]	61.82[a]	64.69[a]	67.05[a]
	(22.9)	(23.35)	(21.03)	(21.53)	(21.13)	(21.49)
MAJPARTY	36.18	33.43	53.91[a]	52.73[b]	68.61[a]	67.0[a]
	(24.48)	(25.19)	(21.54)	(22.34)	(21.50)	(22.1)
TENURE	7.37[a]	7.47[a]	6.86[a]	6.92[a]	6.38[a]	6.41[a]
	(1.93)	(1.99)	(1.85)	(1.91)	(1.82)	(1.84)
GENDER	27.39	29.94	29.47	31.35	30.27	30.83
	(33.99)	(34.44)	(31.89)	(32.61)	(27.19)	(29.01)
HHIE	−.244	.226	−.176	.468	−.125	.393
	(.363)	(.466)	(.338)	(.444)	(.323)	(.413)
PCTBLACK	63.51	3.68	7.43	−28.52	−40.01	−84.72
	(72.22)	(95.88)	(68.7)	(89.26)	(65.73)	(82.19)
PCTCOLLG	312	219	258.3	74.58	195.4	58.91
	(373)	(444)	(354.2)	(426.6)	(338.3)	(398.9)
Ln(*MEDINCM*)	82.4	35.11	95.29	38.81	108.9	59.7
	(75.4)	(99)	(72.77)	(97.52)	(69.5)	(94.9)
1994-Dummy	−36.43[b]	−39.16[b]	−14.38	−19.67	−28.02[c]	−31.69[c]
	(17.8)	(18.43)	(17.24)	(17.67)	(17.02)	(17.2)
1996-Dummy	126.9[a]	128.55[a]	132.2[a]	130.7[a]	113.3[a]	113.5[a]
	(33.15)	(34.49)	(32.6)	(33.95)	(32.93)	(34.0)
1998-Dummy	−59.58[a]	−64.17[a]	−38.17[c]	−44.25[b]	−49.8[a]	−54.9[a]
	(23.09)	(22.38)	(22.74)	(21.57)	(23.41)	(22.46)
Constant	−299	330	−489	206	−662	−56.1
	(754)	(1026)	(709)	(995)	(684)	(969)
State Dummies?	No	Yes	No	Yes	No	Yes
F =	(13,434)=23.30	(55,434)=44.79	(13,434)=27.58	(55,434)=9.38	(13,434)=29.47	(55,434)=39.91
Prob > F =	0.00	0.00	0.00	0.00	0.00	0.00
R^2 =	0.136	0.177	0.171	0.209	0.189	0.225

White's standard errors, in parentheses, are reported to control for variance dependence across groups.
Observations have been clustered by congressional district to control for variance dependence within groups.
Sample is House incumbents seeking re-election in 1992, 1994, 1996, and 1998 (N = 1511).
See Table 1 for definitions, sources, and descriptive statistics of all variables.
a. Significant at the 99 percent confidence level.
b. Significant at the 95 percent confidence level.
c. Significant at the 90 percent confidence level.

majority party and having greater tenure both tend to increase incumbent spending. From the year dummy estimates, it is shown that spending is higher in presidential election years than in midterm elections. Finally, of primary interest is the estimate on *ABSRESID*, which is positive and significant with 99 percent confidence in all of our estimations. Combining the descriptive statistics on *ABSRESID* with the point estimates from Table 4, we can interpret magnitude. Voting one standard deviation (0.529) away from the median voter leads to an estimated increase in incumbent spending between $96,800 (= 51.21/0.529 using Model 1) and $126,750 (= 67.05X0.529 using Model 6).[9] Thus, the evidence available from the 1992 through 1998 elections supports our assertion that incumbents increase expenditures the farther they deviate from their median constituent's interests.[10] This, in turn, suggests that campaign expenditures are persuasive rather than informative.

These results are robust to changes in specification at each stage of the analysis. Slightly changing the set of independent variables in Table 1 (when generating the ADA residuals) does not affect the substantive performance of the residuals in Table 4. Whether year and state dummies are used at any stage alters the point estimates only slightly. Adding and removing individual instrumental variables from Table 3 tends to affect the R^2 in those estimations, and as a result the endogeneity tests perform somewhat differently. But considering how close the 2SLS estimates are to the OLS estimates, this may be inconsequential. Moreover, our variable of primary interest, *ABSRESID*, is positive and significant no matter which estimator is used and no matter which specification is chosen. The preponderance of our evidence indicates that incumbents who deviate more also spend more in their subsequent re-election bid.

CONCLUDING REMARKS

In their survey of the legislator shirking literature, Bender and Lott [1996] point to four areas of relative consensus regarding legislator voting: (a) legislators almost always represent their constituents' interests; (b) when legislators do diverge from constituent interests, the adverse economic effects on constituents are trivial; (c) when legislators do not attempt re-election, their attendance rates fall; and (d) even small deviations from constituent interests quickly lead incumbents to lose re-election. In this study, we continue from this last result by investigating whether incumbents who deviate more frequently or further from their constituents' interests tend to increase their total campaign expenditure to retain their seats, other things equal. Using data from four recent congressional elections, we measure deviation in the usual fashion of the congressional voting literature—as the residuals from a regression of constituent characteristics on the legislator's vote record. We then use the residuals as an independent variable in a model explaining incumbents' re-election campaign expenditures. After controlling for endogeneity with opponent spending and victory margin in a typical model of campaign spending, we find the vote score residual's coefficient to be positive and significant. This suggests that incumbents who deviate more from the median constituent tend to attempt to "buy back" some of the electoral losses from such deviation, providing evidence that such campaign expenditures are persuasive rather than informative, as in the model by Mueller and Stratmann [1994].

Our results are noted with the fact that our current measure of the median voter's position relies on decennial Census data. Many of the Census variables used in this study do not vary over the included sample, so we are limited to cross-sectional estimation techniques. In future work, it would be desirable to measure the median voter's position, and legislators' deviation, using annual data so that longitudinal estimation techniques can be used.

NOTES

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1. Wright [1993] finds that legislators who diverge from constituent preferences lose an average of five percentage points in political support as indicated by primary elections. Lott and Davis [1992] find a significant correlation between shirking and defeat in subsequent election, and Lott and Bronars [1993] show that the House members who lost their re-election bids shirked more in the prior term as a group than those who won re-election. Finally, Kau and Rubin [1993] argue that insofar as ideological shirking exists, it is costly and punished quickly by the electoral process. For an extensive, critical survey of the shirking literature, see Bender and Lott [1996].
2. Note that this resembles two-staged least squares, but the residuals $\hat{u}$ are used in equation (2b) rather than the predicted values for V.
3. Other criticisms are discussed with detail in Bender and Lott [1996].
4. One reader suggested our sample may systematically exclude observations, thus introducing sample selection bias. This can take two possible forms, either by excluding challengers or by excluding incumbents who did not run for re-election. Only incumbents know the true value of in-office deviation from constituent interests and have accurate knowledge of the appropriate price to pay for it. Likewise, only incumbents are able to amass a voting record through vote indices, an opportunity denied to challengers. Still, challengers are effectively included in the model through the variable CHSPEND, discussed presently in the text. Evidence from VanBeek [1991] and Lott [1987] suggests retiring members do not exhibit systematically different voting records, although they do vote less frequently. Based on this evidence, we do not expect significant bias from our selected sample. In any case, there is no effective way to incorporate retirees into our model of re-election. Their vote score residuals during their last period of office go unexplained by our present model. See Lott [1990], Zupan [1990], and Carey [1994] for empirical analyses of shirking and the last period problem.
5. For example, Representative Herb Callahan (Alabama, District 1) shows a raw ADA score of 5 in 1992. For the House, $a_{1992} = 7.27$ and $b_{1992} = .97$, so the adjusted ADA score is −2.34. This reflects a liberal shift of the scale in 1992 relative to other years, so that a score of 5 understates the conservativeness of Rep. Callahan's 1992 vote record, which is more accurately reflected in the score of −2.34. An up-to-date list of a_t and b_t estimates is provided on Tim Groseclose's website at Stanford University (http://faculty-gsb.stanford.edu/groseclose/archive.htm at the time of this writing).
6. Files used in this study include cansum92.zip, cansum94.zip, cansum96.zip, and cansum98.zip.
7. Interpreting the sign of the residuals' true value would be difficult. A positive and significant coefficient on the residuals' true values would suggest that members who deviated more to the right expended more in the next campaign. A negative coefficient would indicate the reverse. We do not provide a theoretical context for either of these results, as our main concern is whether legislators "pay" for deviating from constituents, whether right or left—a concept measured by the residuals' absolute value.
8. To conduct the Hausman test we constructed instruments for *CHSPEND* and *VOTEMARG* as seen in Table 3. We then estimated an OLS model of incumbent spending including the actual data as well as fitted values. A non-zero coefficient estimate on an instrument would indicate correlation with the

OLS residuals, suggesting endogeneity of the variable being instrumented. The estimate on *CHSPEND* is positive and significant with 99 percent confidence, but the estimate on *VOTEMARG* is negative with 87 percent confidence. Considering that the campaign finance literature typically treats *VOTEMARG* as endogenous, we wanted to handle our somewhat borderline result cautiously. In the treatment that follows, we treat both challenger spending and margin of victory as endogenous, but we also obtain results treating only *CHSPEND* as endogenous for comparison. See Hausman [1978] for explanation of the test.

9. Our main result indicates that vote records (V_{kt} as measured by ADA scores or other vote indices) and campaign expenditures are endogenously determined—a problem with which the campaign finance literature has long been concerned [Chappell 1982; Stratmann 1991; 1995; Snyder 1992; Bronars and Lott 1997]. Therefore, using V_{kt} to explain E_{kt} would produce biased coefficient estimates on V. In addition, using μ_{kt} to explain Y_{kt} as in equation (2b) has been shown to introduce bias (Jackson and Kingdon 1992). Whether using μ_{kt} to explain E_{kt} as in equation (3b)—our main estimation—would bias the coefficient estimate on μ has not been discussed in the legislator voting or campaign finance literature as far as we can detect. This represents an interesting econometric question that is introduced by our approach, the investigation of which would contribute to the legislator voting and campaign finance literature.

10. The second stage estimated standard errors have not been adjusted for using estimated values, rather than the actual data, as right-hand-side variables. It is possible to obtain asymptotically correct standard errors, but the calculation is complicated and requires that we have an estimate of Cov(r_C, ε) and Cov(r_M, ε). As Bollen, Guilkey and Mroz [1995] have shown, the available Monte Carlo evidence suggests little improvement is to be had by conducting this complex calculation. Bootstrapping is another option, but the improvement to the standard errors is marginal even with this simpler method. Given the magnitude of the t-statistics on our variables of primary interest in Table 4, neither solution would be likely to alter our substantive findings.

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