Electoral Security and the Provision of Constituency Service

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We examine the relationship between legislators' electoral environment and the provision of constituency service in the Texas State Legislature. Using fictitious constituent requests soliciting information on voter registration and a government program, we analyze the relationship between legislators' previous vote share and the probability of legislator response. To account for possible simultaneity bias of constituency service and election results, we employ an instrumental variables approach. In contrast with previous empirical studies, we find that legislators' response rates to constituent requests decreases in their electoral security across a wide range of model specifications that control for legislator-specific characteristics. We also investigate how electoral security affects legislators' provision of legislative public goods and find some suggestive evidence that electoral security increases the number of bills legislators author, but has little effect on other measures of legislative production.

Motivation

Elections are the fundamental mechanism that citizens use to hold legislators accountable for their actions and, accordingly, electoral incentives play a major role in structuring the behavior of legislators (Mayhew 1974).[1] Empirical studies have analyzed how electoral incentives affect legislators' roll-call voting (Kousser, Lewis, and Masket 2007), appropriations outcomes (Shepsle et al. 2009), and other features of the legislative process (Ragsdale and Cook 1987; Schiller 1995; Sulkin 2005).

In this study we examine how electoral incentives affect the provision of constituency service. We specifically examine how the underlying district electoral environment affects the probability that a legislator responds to a number of constituent requests related to voter registration and distributive politics issues in the Texas State Legislature during September and October 2010. Political scientists have recognized constituency service provision as a key legislative function (Eulau and Karps 1977). Representatives devote a significant share of their time and resources to activities such as responding to mail, answering phone calls, sending out newsletters, and managing casework (Cain, Ferejohn, and Fiorina 1987). For some, these duties represent the central component of their job, more important than bill or policy-related responsibilities (Eulau and Karps 1977). Scholars have argued that officeholders purposefully engage in these important activities to bolster their electoral prospects (Cain, Ferejohn, and Fiorina 1987). Explaining heterogeneity in constituency service provision has implications for the literatures on incumbency advantage, legislative organization, the allocation of scarce legislator resources, interbranch relations, and member-constituent interactions.

Previous empirical work has found mixed results on the relationship between constituency service and electoral margins (Fiorina 1981; Freeman and

[1]An online appendix with supplementary materials for this article is available at www.journals.cambridge.org/jop. We will release all legislator and district specific variables, along with information to reproduce results obtained in the legislative public goods section. As a condition of our Human Subjects Research approval (IRB Protocol 19692 at Stanford University), we agreed to maintain the confidentiality of each legislator. Therefore, we are withholding from release the binary legislator response variable. The data and replication code will be made available upon publication at www.stanford.edu/~zacp.

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Richardson 1996; Johannes 1983). For example, Freeman and Richardson (1996) find no evidence of a statistically significant relationship between electoral margins and casework, as measured by a survey of 310 legislators in four states. In contrast, in a comparative study of legislators in the United States and Great Britain, Cain, Ferejohn, and Fiorina (1987) argue that there is strong evidence that legislators alter their constituency service behavior in response to electoral incentives. In their words, "[r]epresentatives undertake constituency service with a variety of motives, but by their own testimony the electoral incentive is significant, especially for those in marginal districts" (1987, 213–14).

As we elaborate below, we believe that measurement problems in the dependent variable and simultaneity bias explain these contradictory findings. Our approach allows for better measurement of the dependent variable by more closely capturing the environment in which legislators field and respond to requests. Sending fictitious constituent requests to elected officeholders and computing legislators' response rates captures real, albeit small, trade-offs that legislators make in their official duties. We then account for simultaneity bias with an instrumental variables strategy that uses district-level election returns for other races as an instrument for legislators' own vote shares.

Theorists have expressed renewed interest in studying the individual provision of constituency service (Ashworth 2005; Ashworth and Bueno de Mesquita 2006). Ashworth (2005) emphasizes that casework is a more valuable tool for influencing voter beliefs than policymaking. Ashworth and Bueno de Mesquita (2006) develop a model of the optimal provision of constituency service in an environment where voters learn about legislators' ability over time. In equilibrium, legislators who represent districts that are more electorally balanced allocate more of their resources toward constituency service instead of policy work. They cite some empirical evidence that exploits intertemporal variation in aggregate electoral security and the aggregate provision of constituency service. We empirically examine the theoretical result using cross-sectional variation in legislators' electoral characteristics.

Most previous work on constituency service surveys legislators and asks them questions about how the legislator allocates personal and staff time and resources to casework (Freeman and Richardson 1996; Johannes 1983). For example, Freeman and Richardson use a mail survey to ask state legislators whether "[c]onstituency service is the most important thing I do in my position" and "[c]onstituency service is an important method of maintaining electoral support" (1996, 53). While these survey methods are a useful first pass at estimating legislator behavior, they contain weaknesses that restrict the ability of an analyst to make valid inferences.

There are at least three reasons why surveys and elite interviews are of limited use in studying the relationship between electoral incentives and the provision of constituency service. Nonresponse may be correlated with unobservables that affect the legislator's provision of casework. For example, legislators in intense reelection campaigns may not take the time to respond to the survey, systematically biasing the responses. Moreover, our true object of interest should be constituency service, not merely the time input that a legislator puts into providing constituency service, which is what most surveys attempt to measure. Consider the case of two legislators, one from a safe district and one from a competitive district, who each have five staff members to allocate between various legislative tasks. The legislator from the competitive district allocates her two highest-quality staffers to addressing constituent requests while the legislator who represents the safe district allocates her two lowest-quality staffers to this function. Even though both legislators devote the same amount of staff hours to constituency service, there is a meaningful difference in behavior across the two legislative offices. A survey instrument cannot uncover these differences. Most importantly, surveys do not force legislators to make real political tradeoffs between constituency service and policy work. While legislators can claim that constituency service is important, this is very different from devoting scarce staff resources to respond to constituent requests. All of these arguments suggest that a survey-based measure of constituency service will be rife with measurement error. As measurement error in a dependent variable increases standard errors, these measurement problems could be one explanation for the absence of robust findings.

We and previous researchers (Fiorina 1981; Johannes 1983) are concerned about potential simultaneity bias in the relationship between a legislators' previous vote share and constituency service. Suppose a legislator from a competitive district invests heavily in constituency service to cultivate a reputation as a competent incumbent. This may increase her vote share in subsequent elections. If a researcher estimated the naive regression of measured constituency service on vote share, this would tend to attenuate the coefficient estimate on legislator vote share toward 0. We account for this potential bias with an instrumental variables strategy that uses district-level election results for seats on the Texas Railroad Commission and state Supreme Court as an instrument for legislators' own vote shares. Under the identifying assumption that state legislators' provision of constituency service does not affect the vote share in these elections, using these election outcomes as instrumental variables will allow us to recover consistent estimates of the parameters on the relationship between legislators' vote share and the provision of constituency service. We provide a formal econometric justification for our instrumental variables approach in the online appendix.

In contrast to previous empirical work, our analysis reveals a robust negative relationship between legislators' response rates to constituent requests and vote share in the previous election. In our regressions that do not instrument for legislators' vote share, we obtain negative coefficient estimates that are statistically insignificant. Once we instrument for legislators' vote share, the magnitudes of the coefficient estimates increase dramatically,[2] and there is a robust, statistically significant, and negative relationship between legislators' vote share and their response rate. In substantive terms, the probability of responding to a request decreases by 7.5 percentage points when moving from a legislator with the 25th percentile vote share value of 0.602 to a legislator with the 50th percentile vote share value of 0.787, controlling for a number of important legislator and district specific factors. We also use this empirical strategy to examine how electoral security affects legislators' behavior *within* the legislature, in the form of authoring, coauthoring, sponsoring, and cosponsoring bills. If legislators allocate their time between particularistic constituency service and producing legislation, then the same electoral security that decreases constituency service should increase legislative activity, as predicted by Ashworth and Bueno de Mesquita (2006). We find that electoral security increases the number of bills that legislators author.

We make three primary contributions. First, we use an improved methodology that reduces measurement error in the dependent variable and better accounts for the simultaneity of constituency service and legislator vote share. Second, we contribute to the recently revived literature on how electoral incentives affect the behavior of government officials. We document a negative, significant, and politically meaningful relationship between legislator electoral security and provision of constituency service, empirically validating the theoretical predictions of Ashworth and Bueno de Mesquita (2006). Third, we demonstrate that incorporating rich individual-level covariates (legislator characteristics in our setting) and accounting for potential simultaneity bias allows the analyst to differentiate between subtle theoretical mechanisms.[3]

Theoretical Setting

We briefly and informally explain the key details and equlibrium predictions of the Ashworth and Bueno de Mesquita (2006) model. The players in the game are a legislator and a voter in a two-period setting. A legislator is faced with an allocation problem where she chooses to divide her resources between providing particularistic constituency service and legislative public goods. In our setting, this problem can be conceptualized as a legislator's choice in allocating her staff members across the distinct responsibilities of responding to constituent requests and writing legislation. After observing measures of the incumbent's performance in office, the voter reelects the incumbent legislator or replaces her with a challenger.

The legislator values reelection, ideological policy outcomes, constituency service, and the total amount of legislative public goods provided in the legislature. Similarly, the voter values ideology, constituency service, and legislative public goods, but may place different weights on these outcomes than the legislator. The voter's information set is limited in two important ways. First, she does not observe the amount of legislative public goods that the legislator produces.[4] Second, she only observes a noisy signal of the legislator's production of constituency service.

The legislator has an unobserved ability parameter that affects her capacity to provide constituency service and the voter makes Bayesian inferences about the value of this parameter after observing the noisy signal of constituency service. In equilibrium, the voter uses a cutoff rule based on the realization of the signal, which induces a probability distribution over

[2]The point estimate on the legislator vote share coefficient in the instrumental variables probit regressions is approximately three times the magnitude of the coefficient estimate in the naive probit regressions.

[3]In a similar sprit, Grose (2010) studies state legislators' abstention decisions on roll-call votes after treating the legislators with different primes about the legislator's probability of being pivotal on the roll call and potential electoral benefits and costs of voting. Butler and Broockman (2011) send voting registration requests to state legislators across the United States and examine how response rates vary across the racial identity of the fictitious constituent. See also Malhotra (2006) for an audit study of constituency service in the U.S. Senate.

[4]Ashworth (2005) derives similar results with the less restrictive assumption that the precision of the voter's signal is smaller for constituency service than for the provision of legislative public goods.

whether the legislator is reelected for a given allocation between constituency service and legislative public goods for the legislator.

To derive comparative statics on the effect of partisan balance on incumbent behavior, Ashworth and Bueno de Mesquita (2006) examine a setting where the pivotal voter in a district is a realization from a Normal distribution with known mean and variance. This framework allows the partisan leaning of a district to be parameterized as the mean of this distribution. Their Proposition 2 states that the provision of constituency service for leftwing legislators is maximized in moderately right-leaning districts. Noting that left-wing districts are likely to have left-wing legislators and right-wing districts are likely to have right-wing legislators leads to the following corollary: "On average, constituency service is increasing in district-level partisan balance" (2006, 174). This is the key prediction that we take to the data.

Background on the Measurement Protocol

The Texas State Legislature is bicameral with 150 representatives and 31 senators. During our study, one seat in the Senate and one seat in the House were vacant so our sample consists of 179 legislators. House members serve two-year terms and senators serve overlapping four-year terms. There are no formal term limits. Each of the 179 state legislators in our sample received a total of six requests via email from September 20, 2010 through October 19, 2010. We sent the requests during business hours on weekdays (excluding Fridays), and each legislator received a constituent request roughly every three business days. The Texas State Legislature was not in session during this period.

We sent each legislator three requests concerning a government program and three requests soliciting information on voter registration for a total of six email requests. The online appendix contains copies of both requests. Each of the requests came from a unique email account.[5] We randomized the order in which each legislator received the six requests.

The method we used to contact legislators is slightly different from the way constituents typically contact their state legislators in Texas. We obtained the personal email addresses of all state legislators from a publicly available online directory[6] and sent requests directly to these addresses. Since these addresses are not displayed on legislators' web pages, many constituents instead fill out an online form letter with fields for name, address, phone number, and message content. According to conversations with the Texas Senate Research Center, these messages are then routed to legislators' offices. Each office determines whether these requests are sent directly to the legislator's personal email address or to a staffer's inbox. While our means of transmitting the requests might shift the average response rate toward 0 relative to using the legislators' web page form, this potential effect should be orthogonal to our independent variables of interest. To allay concerns that some legislators do not use their personal email addresses, we also report results that only include the sample of legislators who respond to at least one request.[7] Despite the loss of statistical power that emerges from discarding data, the results are very similar to the specifications that include the full sample of legislators.

Empirical Method

After collecting and coding a binary indicator for whether the legislator responded to each request, we estimate a series of regressions examining the relationship between response rates and electoral characteristics. We will refer to three main specifications as the baseline model, the instrumental variables approach, and the normal vote model. In our baseline specification we estimate probit models of the form:

$$Pr(Response_{ik}) = \Phi(\alpha_0 + \alpha_1 VoteShare_i + \beta X_i + \gamma X_k)$$

where we have indexed legislators by i and requests by k. X_i is a vector of legislator-specific covariates, and X_k is a vector of request-specific characteristics. We include indicators for whether the legislator is running for reelection, a member of the Republican

[5]We recruited participants with the following names for the study: Jake Allen, Edwin Batista, Antonio Figueroa, Maurice Thomas, Jamar Walker, and Connor Williams. With their permission, we created email addresses in their names for all correspondences with state legislators.

[6]The directory is available at http://www.txdirectory.com/online/.

[7]If we believe, alternatively, that our method of delivering requests lowers the probability that all legislators respond to a particular request, then the sign of our main parameter of interest is still estimated consistently. Similarly, our approach estimates the true sign if the probability of missing the request is a decreasing function of legislator own vote share. If the probability of missing the request is an increasing function of legislator own vote share, then our estimates will be biased upward, but the underlying mechanism of legislators investing in constituency service as function of own vote share is consistent with our story.

party, and serves in the state senate. We also include linear and quadratic controls for legislators' age, tenure in the chamber, and the distance between the state capital, Austin, and the geographic centroid of the member's district. For X_k, we include control variables for characteristics of the email request, specifically fixed effects for each constituent name and an indicator for whether the request concerns voter registration. We include a table of summary statistics in the online appendix.

We are estimating a latent variable, $Pr(Response_{ik})$, which is a function of the legislator's investment in constituency service. We can decompose this investment, c_i, into the amount of total resources, R_i, that the legislator has in her endowment and the fraction of her resource endowment she invests in constituency service:

$$c_i \equiv R_i \frac{c_i}{R_i} .$$

It is possible that R_i is a decreasing function of the electoral security of the district and that c_i/R_i is constant in the electoral security of the legislator. Alternatively, R_i may be independent of electoral security and the fraction of investment in constituency service is decreasing in electoral security. Either of these mechanisms would generate a negative coefficient estimate on α_1, however, we believe that the fraction of investment in constituency is driving the results because there is little cross-sectional variation in the formal resources that legislators have at their disposal. Under H.R. 3, each member of the Texas House receives a credit to their operating account of \$13,250 (\$12,250) in each month that the legislature is (not) in session. If House members with prior service have not spent all of their funds from the previous fiscal year, they are permitted to transfer up to \$25,000 to their account for the new fiscal year.[8] While formal resources do not vary much across legislators, it is possible that informal resources do vary. For example, if legislators acquire managerial expertise as their tenure increases it is possible that more experienced legislators will be able to generate more production from an equivalent budget. To capture this potential source of variation in resources, we add controls for legislator tenure to our specifications.

An important feature of these data is that we have multiple observations for each legislator, but the independent variables of primary interest do not vary for a given legislator. Treating the disturbance terms as independent in the presence of unobserved legislator heterogeneity will tend to bias standard error estimates downward (Moulton 1990). To account for the grouped structure of the data when making our statistical inferences, we compute heteroskedasticity robust standard errors that are clustered at the legislator level in all of the specifications.

As discussed in the introduction, if constituency service provision also influences a legislator's vote share, the model specified above will yield inconsistent parameter estimates. To obtain consistent parameter estimates in the presence of simultaneity, we instrument for own district vote share with the results of other statewide elections.[9] We use two elections for Texas statewide office on the November 2008 ballot: the race for seat 3 in the Texas Railroad Commission, the primary state energy regulatory agency in Texas, and chief justice of the Texas Supreme Court. Each of these elections pitted Democratic and Republican candidates against Libertarian candidates. The Republicans won both elections with 52.13 and 53.1% to the Democrats' 44.35 and 43.79% in the Railroad Commission and Supreme Court elections, respectively.[10] To compute the instruments for each district, we calculate the fraction of total votes cast for the Democratic and Republican candidates in both of the elections. For districts represented by a Democratic (Republican) legislator, the values of the two instruments are simply the

[8]From our conversations with Texas state legislative employees, the number of staffers in state capitol offices typically ranges from two to six. Some legislators hire unpaid interns to perform office duties. We collected data on the number of staffers for the sample of 143 legislators in office in 2011 from the Texas Legislative Directory. The mean number of staffers listed in the directory is 2.8 and the standard deviation is 1.

[9]In the online appendix, we provide a formal justification that naive OLS results in inconsistent parameter estimates while the IV estimator is consistent.

[10]In the Supreme Court race, Republican Wallace B. Jefferson earned 53.1% of the vote while Democrat Jim Jordan won 43.8%. Jefferson was first appointed to fill a court vacancy in 2001. In the railroad commissioner's race, Republican Michael L. Williams defeated Democrat Mark Thompson by a 52 to 44 margin. Williams was first elected to the commission in 2000. There is little evidence that either Williams or Thompson have benefited from an incumbency advantage, as both received a smaller share of the overall vote in 2008 than they did in 2002. Additionally, aggregate roll-off relative to ballots cast in the presidential election was very low for both of these elections. There were 8,077,795 votes cast in the presidential election and 7,680,041 in the Railroad Commission and 7,705,677 in the Supreme Court election.

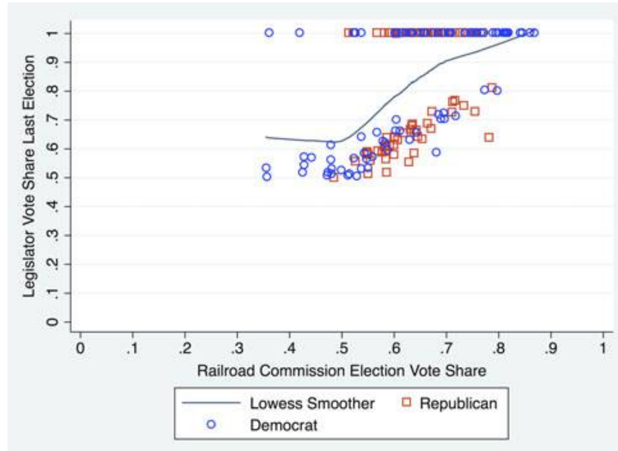
Democratic (Republican) vote share in the Railroad Commission and Supreme Court elections.[11]

In order to be a valid instrument, these variables must satisfy both the inclusion and exclusion restrictions. The inclusion restriction that the instruments are correlated with the endogenous explanatory variables is satisfied at conventional levels. Figure 1 and Figure 2 display scatter plots and loess regressions of the Texas Railroad Commission and State Supreme Court election results against legislators' previous election vote share. The relationship between these variables appears to be fairly strong and monotone. Table 1 reports that the F-statistic of the regression of legislator vote share on the two instruments is approximately 44 in the full sample of responses and 33 when we restrict the sample to legislators who respond to at least one of our requests, so the finite sample problems that emerge from weak instruments (Bound, Jaeger, and Baker 1995) do not pose a serious problem in this setting.

These variables will satisfy the exclusion restriction for a valid instrument if they are uncorrelated with the error term in the response equation. We need vote share in the Texas Railroad Commission and Supreme Court elections to influence a legislator's response probability only through these variables' effect on legislator vote share. One potential mechanism that would induce a violation of the exclusion restriction is if voters in a given district form their beliefs about political parties based on the performance of their state legislator. For example, if a legislator's competent performance in office leads voters in her district to improve her opinions of the Democratic Party, this would lead to similar simultaneity biases to those we seek to avoid through our instrumental variables strategy. However, partisan identity is fairly stable over the life cycle and any residual variation is much more likely to be explained by the well-publicized actions of national political leaders than state legislators.

In our primary specifications, we have chosen to treat legislators' vote share as an endogenous explanatory variable and instrument for this variable using district partisan leanings as expressed in statewide election returns. An alternative approach would be to treat the normal vote as an exogenous variable and simply regress response rates on the district normal vote and other legislator characteristics. We also report results that treat the normal vote—calculated as the average vote share of Democratic (Republican) candidates in the Railroad Commission and Supreme Court election for Democratic (Republican) legislators—as an exogenous independent variable.[12] As we show below, these results are quite similar to our instrumental variables specifications.

Figure 1 Scatter Plot and Nonparametric Regression of Legislator Vote Share on Railroad Commission Election Results



Results

Figure 3 plots a histogram of the number of responses to the six email requests for the 179 legislators in our study. Among the legislators who respond to at least one request the modal number of responses is 3. There are 39 legislators who respond to 0 requests.

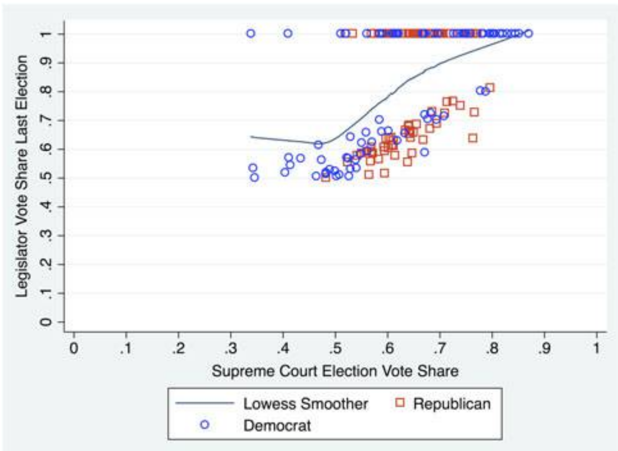
Before proceeding to the multivariate analysis, we examine the bivariate relationship between the legislator's electoral environment and their response rate to the email requests.[13] Figure 4 plots the loess

[11]We also collected data on vote share in the 2008 presidential election in every district. In this relatively high-information setting, idiosyncratic features of the presidential candidates and districts may affect vote shares. We believe that the low-information state election results are more appealing instrumental variables because they tap into the fundamental partisan identification of district voters. However, using presidential vote share as our instrument leads to qualitatively similar results, which are available on request.

[12]Another potential advantage of these specifications is that they allow legislators' responsiveness to vary directly with underlying district electoral characteristics. In contrast, the instrumental variables specifications generate predicted response probabilities that do not vary with district partisan characteristics beyond legislators' own vote share. We thank an anonymous referee for this observation.

[13]The figures only display the nonparametric loess regression curve and not the underlying observations used to compute the curve. When seeking Institutional Review Board approval for this project, the IRB manager requested additional procedures to ensure that legislators' response rates remained anonymous. To comply with this request, we do not display any of the data points and only report aggregate relationships in this article.

FIGURE 2 Scatter Plot and Nonparametric Regression of Legislator Vote Share on Supreme Court Election Results



nonparametric regression of a legislator's response rate to the requests on her vote share in her most recent election.[14] The resulting regression line shows a general downward trend in response rates as legislators have higher vote shares. The highest response rates are in the neighborhood of 0.5 with a moderate and uneven decrease in response rates until a steep descent occurs in the area of 0.9. The nonparametric regressions provide some suggestive evidence that legislators' responses to constituent requests vary according to district electoral characteristics, but they cannot control for additional individual legislator characteristics such as age and experience in the legislature. Further, the statistical power of these regressions is quite limited.

Table 2 displays the results of a series of probit regressions of the binary response indicator on legislator characteristics. These regressions do not account for the potential simultaneity of a legislator's vote share in the previous election. As we discussed above, these regressions do not provide credible causal estimates, but we present them as a baseline for comparison with the subsequent model specifications. In the first three columns, we estimate the models on the full sample of email requests. In all three regressions, the coefficient estimate on legislator vote share is negative, but insignificant at conventional levels. In columns 4–6, we restrict our analysis to the sample of legislators who respond to at least one of our requests. Again, the coefficient estimates

[14]In this and all subsequent empirical analysis, we use two-party vote share instead of nominal vote share.

TABLE 1 First-Stage Results

	(1)		(2)	
	Full Sample		≥ One Response	
Railroad Commission Election Vote Share	2.431**	(0.927)	2.719**	(0.994)
Supreme Court Election Vote Share	-1.369	(0.930)	-1.588	(0.993)
Constant	0.139^{+}	(0.0802)	0.101	(0.0973)
Observations	179		140	
R^2	0.313		0.326	
F	43.91		32.70	

Standard errors are robust to heteroskedasticity.

on legislator vote share are negative and are insignificant at the 95% level.

Table 3 displays coefficient estimates for models that account for the endogeneity by instrumenting for legislator vote share using the vote share instruments described above. In columns 1, 2, and 3 we present linear two-stage least squares results. The coefficient estimates are negative, statistically significant, and politically meaningful. To put the point estimates in perspective, the coefficient estimate of -0.407 in column 3 implies that the the probability of responding to a request decreases by 7.5 percentage points when moving from a legislator with the 25th percentile vote share value of 0.602 to a legislator with the 50th percentile vote share value of 0.787. The magnitude of this effect is larger than the coefficient estimate on the running for reelection variable.

Since we have two instrumental variables for the single endogenous explanatory variable, these models

FIGURE 3 Histogram of Number of Responses to the Six Email Requests Sent to Each Legislator

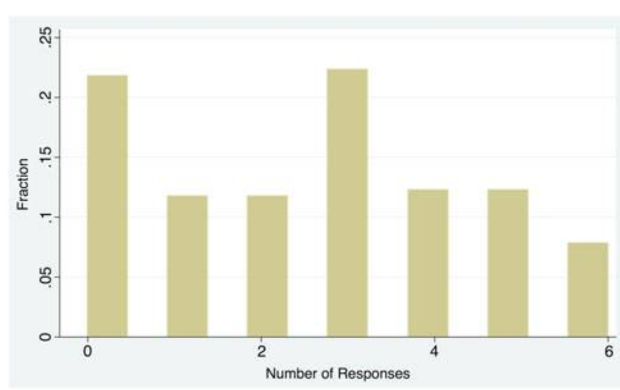
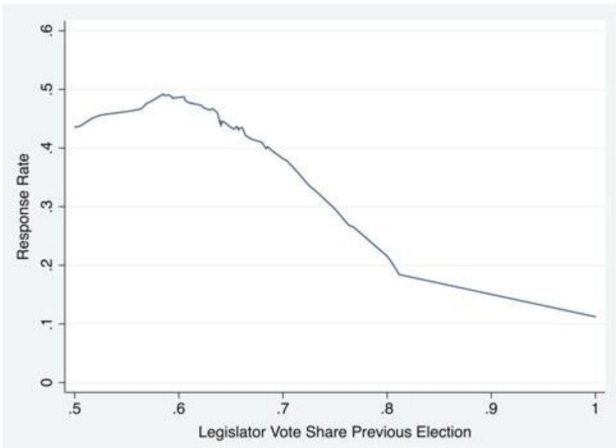


FIGURE 4 Nonparametric Regression of Legislator Response Rates on Legislators' Vote Share in the Most Recent Election Excluding Uncontested Seats



are overidentified. We calculate the Hansen's J statistic (Hansen 1982) to examine whether the models are correctly specified. Under the null hypothesis of a correctly specified model Hansen's J statistic is distributed chi-squared with one degree of freedom. As seen at the bottom of Table 3, in each of the specifications we fail to reject the null at the 90% significance level, providing additional support for our models.

In addition to our primary independent variable of interest, the regression results shed light on other factors that explain variation in constituency service.[15] The coefficient on the state senator indicator variable is uniformly positive across the specifications and marginally statistically significant in some of the specifications. Republican legislators have a significantly higher response probability across specifications. The effects of age and chamber tenure on response rate appear to be quite small. The indicator for whether the legislator is running for reelection is positive, but insignificant in all of the specifications. The distance from Austin variable is estimated to be negative and statistically significant. This evidence is consistent with a story where legislators who represent outlying districts spend more time traveling and less time on legislative work.[16]

In columns 4, 5, and 6 of Table 3, we report results conditioning on the sample of legislators who respond to at least one of the requests. The primary difference between the results in the conditioned sample is that the legislator vote share variable is insignificant in column 4 of Table 3. Once we include the additional legislator covariates, this coefficient estimate is significant at the 95% level.[17]

While these linear probability models are easy to interpret and are preferred by many applied researchers as results, they do not always predict response probabilities in the [0, 1] interval. To account for the binary nature of the dependent variable, we also estimated instrumental variables probit regressions. Comparing the results in Table 2 to Table 4, we see that the point estimates are dramatically attenuated toward 0 in the former. This could explain the absence of robust findings in previous empirical studies on the relationship between constituency service and the electoral environment that failed to account for the endogeneity of a legislator's vote share. In the final three columns, we again condition the analysis on the sample of legislators who respond to at least one request and obtain similar results.

We also estimate probit regressions where the independent variable of interest is the district normal vote, as measured by the average two-party vote share of the legislators' copartisan candidates in the Supreme Court and Railroad Commission elections. As seen in Table 5 the results are qualitatively similar to

[15]In unreported results, we include the number of staffers for the sample of 143 legislators for which we observe this variable. The coefficient on the number of staffers is not significantly different from 0 and including this variable does not qualitatively affect the results.

[16]Additionally, the coefficient estimate on the quadratic distance term is estimated to be positive. This nonmonotone relationship between responsiveness and distance would emerge in the data if moderately distant legislators travel to Austin by automobile and legislators who represent more distant districts travel by airplane.

[17]As an additional robustness check, we bootstrap the standard errors in our instrumental variables specifications. Although the standard errors that we report in all of our specifications are clustered at the legislator level, it remains possible that there are complex interdependencies in the error structure that could lead to underestimating the standard errors. To account for this possibility, we use the nonparametric block bootstrap with 200 replications. In an online appendix, we report these results for the full sample and the sample of legislators who respond to at least one request. The standard errors on legislator vote share are slightly larger than those calculated from clustering, but this does not change the results of our hypothesis tests at conventional significance levels.

TABLE 2 Naive Probit Regressions of Response on Previous Legislator Vote Share and Controls

	(1)		(2)		(3)		(4)		(5)		(6)	
	Full Sample						>=One Response					
Legislator Previous Vote Share	-0.210	(0.321)	-0.180	(0.337)	-0.212	(0.328)	-0.183	(0.290)	-0.202	(0.312)	-0.266	(0.304)
Running for Reelection	0.169	(0.204)	0.158	(0.216)	0.0519	(0.211)	-0.0200	(0.188)	-0.0283	(0.206)	-0.0474	(0.197)
Senate	0.285	(0.189)	0.381^{+}	(0.198)	0.372^{+}	(0.193)	0.138	(0.179)	0.174	(0.192)	0.205	(0.183)
Chamber Tenure			-0.0450	(0.0286)	-0.0418	(0.0274)			-0.0531*	(0.0255)	-0.0505*	(0.0246)
Chamber Tenure^2			0.00133	(0.00104)	0.00143	(0.00100)			0.00222**	(0.000843)		
Age			-0.0893^{+}	(0.0517)	-0.139**	(0.0511)			-0.0525	(0.0494)	-0.0710	(0.0508)
Age^2			0.000857^{+}	(0.000469)	0.00126**	(0.000461)			0.000474	(0.000452)	0.000639	(0.000462)
Republican					0.417**	(0.138)					0.169	(0.128)
Miles from Austin (100s)					-0.00525**	(0.00177)					-0.00497**	(0.00155)
Miles From Austin^2 (100s)					0.00000637^{+}	(0.00000356)					0.00000978***	(0.00000291)
Constant	-0.511	(0.356)	1.926	(1.482)	3.890**	(1.489)	-0.105	(0.337)	1.505	(1.393)	2.438^{+}	(1.449)
Name and Request Issue FEs	Yes		Yes		Yes		Yes		Yes		Yes	
Observations	1074		1074		1074		840		840		840	
Pseudo R^2	0.019		0.034		0.069		0.025		0.036		0.049	

Standard errors in parentheses are robust to heteroskedasticity and clustered at the legislator level. The first three columns are models estimated on the entire sample while columns 4-6 only include legislators who responded to at least one request.
$+$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

TABLE 3 Instrumental Variables Linear GMM Regressions

	(1)		(2)		(3)		(4)		(5)		(6)	
	IV-Linear GMM						>=One Response					
LegVSTwoParty	-0.465*	(0.223)	-0.545*	(0.223)	-0.407^{+}	(0.213)	-0.263	(0.198)	-0.434*	(0.208)	-0.446*	(0.199)
Running for Reelection	0.0550	(0.0771)	0.0708	(0.0790)	0.0191	(0.0758)	-0.0177	(0.0690)	-0.00124	(0.0741)	-0.0200	(0.0700)
Senate	0.114	(0.0700)	0.161*	(0.0707)	0.136*	(0.0670)	0.0494	(0.0658)	0.0813	(0.0684)	0.0858	(0.0644)
Voting Issue	0.118***	(0.0279)	0.118***	(0.0278)	0.113***	(0.0277)	0.147***	(0.0352)	0.147***	(0.0352)	0.146***	(0.0351)
Chamber Tenure			-0.00866	(0.0107)	-0.00907	(0.00968)			-0.00906	(0.00810)	-0.00840	(0.00791)
Chamber Tenure^2			0.000283	(0.000366)	0.000390	(0.000337)			0.000434*	(0.000201)	0.000429*	(0.000199)
Age			-0.0449*	(0.0216)	-0.0603**	(0.0198)			-0.0355^{+}	(0.0203)	-0.0404*	(0.0203)
Age^2			0.000426*	(0.000194)	0.000544**	(0.000178)			0.000323^{+}	(0.000187)	0.000363^{+}	(0.000186)
Republican					0.165***	(0.0482)					0.0821^{+}	(0.0482)
Miles from Austin (100)					-0.195**	(0.0643)					-0.193***	(0.0560)
Miles from Austin^2 (100)					0.0243^{+}	(0.0133)					0.0393***	(0.0103)
Constant	0.627**	(0.207)	1.845**	(0.668)	2.416***	(0.606)	0.625***	(0.185)	1.707**	(0.614)	2.007***	(0.601)
Name and Request Issue FEs	Yes		Yes		Yes		Yes		Yes		Yes	
Observations	1074		1074		1074		840		840		840	
Hansen's J statistic	0.0813		0.448		2.611		0.134		0.00835		0.632	

This table reports over-identified instrumental variables regressions where we instrument for legislator's vote share using the district vote share of the Democratic (Republican) candidates for the Texas Railroad Commission and Supreme Court chief justice for Democratic (Republican) legislators.

Standard errors in parentheses are robust to heteroskedasticity and clustered at the legislator level.

$+$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table 4 Instrumental Variables Probit Regressions

	(1)		(2)		(3)		(4)		(5)		(6)	
	Full Sample						>=One Response					
Legislator Previous Vote Share	-1.181*	(0.559)	-1.409*	(0.565)	-1.146*	(0.571)	-0.672	(0.509)	-1.112*	(0.528)	-1.164*	(0.513)
Running for Reelection	0.138	(0.202)	0.177	(0.208)	0.0680	(0.210)	-0.0357	(0.184)	-0.00933	(0.196)	-0.0239	(0.191)
Senate	0.283	(0.179)	0.396*	(0.183)	0.385*	(0.185)	0.131	(0.173)	0.187	(0.184)	0.223	(0.176)
Voting Issue	0.295***	(0.0710)	0.297***	(0.0714)	0.313***	(0.0750)	0.375***	(0.0911)	0.376***	(0.0911)	0.383***	(0.0924)
Chamber Tenure			-0.0209	(0.0288)	-0.0235	(0.0275)			-0.0363	(0.0261)	-0.0328	(0.0250)
Chamber Tenure^2			0.000696	(0.00102)	0.000945	(0.000980)			0.00175*	(0.000878)	0.00160$^+$	(0.000838)
Age			-0.119*	(0.0556)	-0.163**	(0.0542)			-0.0894$^+$	(0.0536)	-0.107$^+$	(0.0555)
Age^2			0.00113*	(0.000502)	0.00147**	(0.000487)			0.000810	(0.000495)	0.000963$^+$	(0.000508)
Republican					0.429**	(0.135)					0.203	(0.129)
Miles from Austin (100)					-0.495**	(0.185)					-0.507**	(0.157)
Miles from Austin^2 (100)					0.0587	(0.0388)					0.104***	(0.0295)
Constant	0.324	(0.530)	3.560*	(1.702)	5.144**	(1.653)	0.311	(0.483)	3.110$^+$	(1.600)	3.988*	(1.642)
Name and Request Issue FEs	Yes		Yes		Yes		Yes		Yes		Yes	
Observations	1074		1074		1074		840		840		840	
Pseudo log-likelihood	-293.8		-238.6		-201.3		-230.4		-177.9		-158.2	

This table reports over-identified instrumental variables regressions where we instrument for legislator's vote share using the district vote share of the Democratic (Republican) candidates for the Texas Railroad Commission and Supreme Court chief justice for Democratic (Republican) legislators.
Standard errors in parentheses are robust to heteroskedasticity and clustered at the legislator level.
$+$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

TABLE 5 Probit Regressions with Normal Vote as Exogenous Independent Variable

	(1)		(2)		(3)		(4)		(5)		(6)	
	Full Sample						≥ One Response					
Normal Vote	-1.256*	(0.619)	-1.452*	(0.609)	-1.522*	(0.623)	-0.812	(0.585)	-1.207*	(0.586)	-1.394*	(0.562)
Running for Reelection	0.142	(0.201)	0.138	(0.211)	0.0359	(0.204)	-0.0269	(0.186)	-0.0342	(0.203)	-0.0554	(0.191)
Senate	0.263	(0.188)	0.357^{+}	(0.196)	0.348^{+}	(0.193)	0.137	(0.178)	0.176	(0.192)	0.209	(0.185)
Chamber Tenure			-0.0441	(0.0276)	-0.0414	(0.0264)			-0.0558*	(0.0244)	-0.0540*	(0.0236)
Chamber Tenure^2			0.00147	(0.00102)	0.00160	(0.000983)			0.00243**	(0.000852)	0.00236**	(0.000854)
Age			-0.0963^{+}	(0.0505)	-0.146**	(0.0499)			-0.0637	(0.0476)	-0.0845^{+}	(0.0503)
Age^2			0.000919*	(0.000457)	0.00132**	(0.000448)			0.000570	(0.000436)	0.000751	(0.000457)
Republican					0.439**	(0.135)					0.199	(0.126)
Miles from Austin (100)					-0.539**	(0.176)					-0.514***	(0.152)
Miles from Austin^2 (100)					0.0687^{+}	(0.0364)					0.102***	(0.0288)
Constant	0.146	(0.459)	2.891^{+}	(1.493)	4.874**	(1.504)	0.265	(0.432)	2.423^{+}	(1.392)	3.495*	(1.486)
Name and Request Issue FEs	Yes		Yes		Yes		Yes		Yes		Yes	
Observations	1074		1074		1074		840		840		840	
Pseudo R^2	0.026		0.042		0.077		0.027		0.042		0.056	

This table reports normal vote regressions where we compute district preferences as the average of the vote share of Democratic (Republican) candidates for the Texas Railroad Commission and Supreme Court chief justice for Democratic (Republican) legislators.
Standard errors in parentheses are robust to heteroskedasticity and clustered at the legislator level.
$+$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

TABLE 6 Legislative Production Regressions, IV-Linear GMM

	(1)		(2)		(3)		(4)	
	Authored		Co-authored		Sponsored		Co-sponsored	
Legislator Previous Vote Share	65.31*	(33.09)	0.210	(11.37)	-6.593	(5.207)	-1.480	(1.290)
Running for Reelection	6.015	(11.61)	-5.309	(4.812)	7.265*	(2.947)	-0.217	(0.757)
Senate	54.98***	(13.13)	-2.602	(4.627)	32.86***	(4.266)	1.994**	(0.712)
Chamber Tenure	1.913	(1.295)	-1.247*	(0.529)	0.616*	(0.287)	-0.116*	(0.0593)
Chamber Tenure^2	-0.0218	(0.0335)	0.00686	(0.0136)	-0.0150$^+$	(0.00837)	0.00291$^+$	(0.00174)
Age	0.745	(2.606)	-1.435	(1.071)	-0.495	(0.497)	-0.0118	(0.0824)
Age^2	-0.0104	(0.0237)	0.0147	(0.00987)	0.00339	(0.00446)	0.000175	(0.000742)
Republican	-27.56***	(8.199)	-1.055	(3.164)	1.246	(1.567)	-1.533***	(0.394)
Miles from Austin (100)	-18.63$^+$	(10.19)	-2.858	(3.606)	-0.703	(2.529)	0.495	(0.506)
Miles from Austin^2 (100)	3.675*	(1.781)	0.421	(0.614)	0.295	(0.454)	-0.0739	(0.0755)
Constant	25.03	(76.28)	87.12**	(29.78)	19.58	(15.72)	3.547	(2.807)
Hansen's J statistic	2.507		0.760		0.137		0.636	

Standard errors in parentheses are robust to heteroskedasticity.
+ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

those that instrument for legislator vote share. Again, we see that the coefficient estimates in Table 5 are negative and statistically and politically significant.[18] The coefficient estimates are on the order of five times the magnitude of those reported in Table 2. This comparison vividly illustrates how simultaneity attenuates the magnitude of the coefficient estimates in the naive regression.[19]

[18]We also estimated models with quadratic terms for the normal vote variable. While we have emphasized the prediction of a monotone decreasing relationship between electoral security and the provision of constituency service, Ashworth and Bueno de Mesquita (2006) show that Democratic (Republican) legislators who represent strongly right-leaning (left-leaning) districts will provide lower levels of constituency service than legislators who represent moderately right-leaning (left-leaning) districts. In our estimation, we do not find a significant quadratic term. Our interpretation of this empirical result is that there are insufficiently many Democratic (Republican) legislators who represent districts that are strongly Republican (Democratic) to generate a nonmonotone relationship in the data. An additional prediction of Ashworth and Bueno de Mesquita (2006) is that these extremely out of step legislators will have low reelection probabilities so our empirical finding is reconcilable with the theoretical framework. The regression results are available on request.

[19]We also separately estimated the instrumental variables linear, instrumental variables probit, and normal vote specifications with the full battery of controls for Democratic and Republican legislators on the full sample and the sample of legislators who respond to at least one request. Although partitioning the sample reduces the statistical power of the analysis, the results are fairly robust. Each of the six coefficient estimates from the sample of Democratic legislators is negative and statistically significant at the 90% level. Each of the six coefficient estimates from the sample of Republican is negative and three are statistically significant at the 90% level. The three remaining p-values are 0.119, 0.167, and 0.204. The full set of results are available upon request.

Legislative Public Goods

If legislators allocate their resources between particularistic constituency service and the production of legislation, we should expect that legislators' provision of legislative public goods is lower for marginal legislators compared with their electorally secure counterparts.[20] A crude measure of an individual's production of legislative public goods in the Texas state legislature is the number of bills that they author, coauthor, sponsor and cosponsor.[21] We collected these data for the 81st regular session for each member.[22]

[20]In addition to providing constituency service and legislative public goods, legislators may shirk in their legislative duties by allocating their time to private production or leisure. If we introduce this additional option for the legislators and assume that voters' do not value legislative public goods, then the relationship between electoral security and the production of legislative public goods will not necessarily vary with electoral security and will instead depend on legislator-specific taste parameters for legislative public goods.

[21]In their work on legislative effectiveness in the U.S Congress, Volden and Wiseman (2009) examine similar measures and also how far in the legislative process—such as whether a bill comes up for a floor vote, is passed by the chamber, or becomes public law—different bills advance.

[22]This evidence is much more speculative than our previous findings. A number of factors limit the interpretation of these results. Most importantly, our measure of legislative public good provision is inferior to our measure of constituency service. These legislative activities may serve multiple purposes including signaling to the electorate. Another important obstacle in interpreting these results is that legislators' access to producing legislation may systematically vary over the course of her career and across legislators in unobservable ways Padro i Miquel and Snyder Jr. (2006). Nonetheless, we believe it is useful to examine this relationship and compare the relationship between electoral security across the different measures of legislative production.

Table 6 reports the results of linear instrumental variables regressions of the four different measures of legislative production on all of the legislator covariates we have previously used. Again, we instrument for a legislator's previous vote share using the Texas Railroad Commission and Supreme Court election results. The value of Hansen's J statistic is also insignificant so we again do not reject the validity of any of the specifications. The only coefficient specification where the coefficient estimate on legislator vote share is significant at the 10% level is the specification with bills authored as the dependent variable.[23] As predicted by the theory, the coefficient estimate is positive. The coefficient estimates are negative but insignificant at conventional levels for the three other dependent variables. The null result on the relationship between cosponsorship and electoral security is consistent with the findings of Kessler and Krehbiel (1996) in their study of cosponsorship dynamics in the U.S. House and corroborates their argument that cosponsorship's primary purpose is for intralegislative signalling and not election-oriented position taking.

Conclusion

We have studied the relationship between legislators' electoral characteristics and their provision of constituency service. Our key innovation is the use of a new measurement technique that forces legislators to make real trade-offs in responding to constituency requests, as opposed to survey- and interview-based measures of legislators' constituency service used in previous studies (Freeman and Richardson 1996; Johannes 1983). In contrast to much previous empirical work and in support of the predictions of Ashworth and Bueno de Mesquita (2006), we have found robust evidence that higher levels of electoral safety tend to decrease legislators' supply of constituency service. We have also pointed to some more speculative evidence that the production of legislative public goods, specifically the number of bills authored, increases with electoral security.

While the analysis is restricted to a single legislature, the state of Texas is important in its own right. As the second largest state in the Union, each state Senator represents about 800,000 constituents and each state House member represents approximately 165,000 Texans.[24] These elected officeholders are responsible for debating, modifying, and approving a biennial budget that approaches $200 billion. Further, our results are relevant beyond the legislative chambers in Austin. State legislatures constitute training grounds for future leaders, such as governors and federal legislators. The organizational skills a legislator develops in this setting shape her responsiveness in future offices.

In spite of our project's limited geographic reach, the use of Texas as the subject of this study is unlikely to lead to biased or incomparable findings. For instance, Texas is near the median on a prominent legislative professionalism index that uses legislator compensation, calendar length of sessions, and staff resources to compare states (Squire, 2007).[25] Selecting a state with unique institutional features—such as a single legislative chamber or multi-member districts—or choosing a case at either extreme of the aforementioned legislative professionalism index would call into question the generalizability of the findings.

Our analysis has focused on underlying district political preferences that do not vary much over time within district.[26] An alternative analysis could examine how race-specific features of legislator vulnerability affect legislators' provision of constituency service. To take a stark example, consider two legislators who represent similarly competitive districts. One legislator is running unopposed while the other is facing a challenger who has previously held public office. While the latter legislator's performance in office could be the determining factor in whether she secures reelection, our empirical specifications treat these two hypothetical legislators identically. From our analysis we are unable to determine whether the observed negative relationship between district security and the provision of constituency service is mediated by challenger quality, aggregate partisan swings, or some other factor.

Another limitation of our research design is the inability to examine dynamic processes. The cross-sectional nature of our data does not give us any within-legislator variation in the intensity of electoral competitiveness. This feature of the data prevents us from determining whether our empirical findings are the result of competitive districts selecting legislators who have a taste for constituency service or legislators

[23]While the p-value on the legislator vote share coefficient estimate (0.048) is less than the conventional cutoff of 0.05, altering the specification—for example, removing the distance from Austin variables—generates p-values that are slightly above this cutoff. Given the sensitivity of this finding, we are more comfortable claiming that the coefficient estimate is significant at the 10 percent level rather than the 5% level.

[24]This is based on 2009 Census estimates.

[25]On this index, Texas places 15th in the 2003 rankings with a score of 0.199. The overall mean is 0.184 and the median is 0.154 in 2003. The scale ranges from 0 to 1.

[26]We are excluding the possibility of redistricting.

dynamically adjusting their provision of constituency service to the electoral environment. A panel-data analysis that includes multiple observations for the same legislators over time could exploit time-series variation in legislators' electoral security to adjudicate between these two hypotheses. Separating out these alternative mechanisms and quantifying their effects on legislators' provision of constituency service is an important avenue for future research.

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