Responsiveness and Election Proximity in the United States Senate

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Abstract: One of the most important questions in the study of democratic representation is whether elected officials are responsive to the preferences of their constituents, and whether responsiveness varies across institutional conditions. However, previous work on this question has been hampered by the unavailability of time-varying data on public opinion in each constituency. In this paper, I use new estimates of public opinion in each state-year from 1950-2012 to examine whether Senators are responsive to changes in public opinion and whether their behavior shifts over the course of the electoral cycle. I find that Senators are modestly responsive to changes in public opinion. They are particularly responsive to public opinion in the last two years of their term. But the impact of public opinion on Senators' roll call behavior is still small relative to the impact of electoral selection. This analysis resolves earlier ambiguities in the literature on election proximity in the Senate, and opens up new research paths in the study of representation.

One of the most important questions in the study of democratic representation is whether elected officials are responsive to the preferences of their constituents. There are strong theoretical reasons to believe that legislators should be responsive to their constituents in order to enhance their re-election prospects (Downs, 1957; Mayhew, 1974; Arnold, 1992). They should be particularly responsive to *changes* in constituent preferences (Stimson, MacKuen, and Erikson, 1995). Moreover, there are a variety of reasons to believe that legislators should be most responsive to public opinion toward the end of the electoral cycle. Most importantly, voters tend to overweight recent events (Achen and Bartels, 2004; Huber, Hill, and Lenz, 2012; Healy and Lenz, 2014). This is likely to incentivize politicians to consider the views of their constituents more as elections grow more proximate.

There is a clear consensus among scholars that more liberal constituencies tend to elect legislators that take liberal roll call positions (Miller and Stokes, 1963; Erikson and Wright, 2000; Ansolabehere, Snyder, and Stewart, 2001; Clinton, 2006). However, there is an active debate about whether legislators respond to changes in the preferences in their constituents. One set of studies finds no evidence that legislators shift their roll call behavior (Poole and Rosenthal, 2000; Lee, Moretti, and Butler, 2004; Krimmel, Lax, and Phillips, 2015). Rather, they "die in their ideological boots" (Poole, 2007, 435). Another set of studies finds some evidence of movement in response to large, exogenous shifts in constituency opinion, such as redistricting (Stratmann, 2000), variation in turnout due to rainfall (Henderson and Brooks, Forthcoming), moves between chambers (Miler, Forthcoming), or during statewide crises (Kousser, Lewis, and Masket, 2007). However, previous research on responsiveness has been hampered by the lack of time-varying measures of citizens' policy preferences in

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each constituency.[1] Only with time-varying measures is it possible to trace out the dynamic process of responsiveness and identify how legislators respond to changes in public opinion.

A number of previous studies have examined how senators' roll call behavior shifts over the electoral cycle. These studies have found that Senators become more moderate as elections approach, presumably because they are shifting their behavior toward the median voter in their state (Albouy, 2011; Amacher and Boyes, 1978; Elling, 1982; Lindstädt and Van-der Wielen, 2011; Thomas, 1985; Wright and Berkman, 1986).[2] But these studies too suffer from the "lack ... [of] a measure that taps actual constituency views against which to compare senatorial voting records" (Ahuja, 1994). The unavailability of time-varying measures of public opinion in each state has forced most previous studies to focus on the average change in roll call behavior over the electoral cycle without reference to a measure of constituency preferences. The previous studies on how responsiveness to public opinion varies over the electoral cycle use cross-sectional data from just one or two election cycles, and thus lack a credible causal identification strategy (Ahuja, 1994; Shapiro et al., 1990).

In this paper, I use new data and statistical techniques to examine democratic responsiveness and election proximity in the United States Senate. First, I examine dynamic responsiveness using new estimates of citizen policy conservatism for each state and year that are derived from the survey responses of over 700,000 individuals between 1950 and 2012. To summarize this large body of opinion data, I estimate a dynamic group-level item-response model, producing yearly estimates of the average economic policy conservatism of

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[1] Only three previous studies have examined whether senators respond to changes in constituent preferences (Levitt, 1996; Wood and Andersson, 1998; Buttice and Highton, Forthcoming). However, each of these studies uses proxies for constituent preferences rather than direct measure of public opinion.

[2] In a related study, Huber and Gordon (2004) find that elected judges change their behavior in criminal sentencing cases as their reelection approaches.

the citizens of each state (Caughey and Warshaw, 2015). To measure legislators' roll call behavior, I use two different dynamic ideal point models to estimate how Senators' ideal points change over time (Martin and Quinn, 2002; Nokken and Poole, 2004). With these new data in hand, I test whether Senators are responsive to changes in the views of their constituents. I find that Senators are modestly responsive to shifts in public opinion: when the state public moves to the left, Senators takes more liberal roll call positions. Moreover, Senators are substantially more responsive to public opinion in the last two years of their term than in the first four years of their term.

As a robustness check for my primary results, I also examine how the association between senators' roll call votes and public opinion on individual issues changes over the course of the electoral cycle. This approach follows that of studies such as Gilens (2005), Lax and Phillips (2009a), Matsusaka (2010), and Krimmel, Lax, and Phillips (2015). One benefit of focusing on the link between individual roll call votes and issue-specific public opinion is that this approach makes fewer assumptions about the dimensionality of the roll call agenda and the public's policy preferences (Broockman, 2016).[3] It also makes it possible to compare public opinion and roll call votes on the same scale (Lax and Phillips, 2009a; Matsusaka, 2010). I examine the association between Senators' roll call votes and public opinion on individual issues by using multilevel regression with post-stratification (MRP) models to estimate issue-specific public opinion on 72 individual policies, and then linking public opinion on each policy directly to senators' roll call votes. This analysis also indicates that there is a much stronger association between Senators' roll call votes and public opinion

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[3] However, while this approach is well-suited for examining how the association between public opinion and roll call behavior varies over the electoral cycle, it is not suitable for causally identifying the degree of Senators' responsiveness to public opinion.

in the last two years of their term.

Overall, my results establish a clear link between public opinion and senators' roll call votes. Moreover, legislators are most responsive to their constituents toward the end of the electoral cycle. However, a one standard deviation change in citizen policy conservatism corresponds to only a 0.02 standard deviation change in Senators' roll call behavior. Even in the last two years of senators' term, a one standard deviation change in citizen policy conservatism only corresponds to roughly a 0.04 standard deviation change in their roll call behavior. So it is true that senators are responsive to changes in public opinion, and more responsive in the last two years of their term, but the aggregate effect of public opinion is still small.

This paper proceeds as follows. First, I describe previous work on the effect of public opinion and election proximity on representation in the Senate, and describe my theory and hypotheses. Then, I describe my research design for measuring dynamic responsiveness. Next, I describe my main findings. Then, I check the robustness of my results using data about public opinion and roll call votes on individual issues. Finally, I briefly conclude and discuss the implications of my findings for our understanding of democratic representation in Congress.

Dynamic Responsiveness

The central assumption of American legislative scholarship over the last 30 years is that legislators are predominately, if not single-mindedly, motivated by electoral incentives (Mayhew, 1974; Kousser, Lewis, and Masket, 2007). Legislators' re-elections hopes are likely to

be bolstered by adapting their positions to respond to changes in the opinion of their constituents. Thus, "when politicians perceive public opinion change, they adapt their behavior to please their constituency, and, accordingly, enhance their chances of reelection" (Stimson, MacKuen, and Erikson, 1995, 545).

There is a broad array of evidence that citizens are capable of holding their legislators accountable for their roll call behavior. Using survey data from the 1988-1992 Senate Election Study, Ensley (2007) shows that the ideological divergence between Senate candidates influences citizens' vote choices. Moreover, Ansolabehere and Jones (2010), Jessee (2009), and Shor and Rogowski (2013) find that the extent to which a constituent agrees with the policy positions of their legislator affects the constituent's likelihood of voting for the member. Looking at electoral results, there is also a wide array of empirical evidence that ideologically extreme legislators are penalized at the ballot box (Ansolabehere, Snyder, and Stewart, 2001; Canes-Wrone, Brady, and Cogan, 2002; Hall, 2015). This leads to the prediction that legislators should be attentive and responsive to changes in the policy preferences of their constituents.

However, several factors may dampen the incentives for legislators to be highly responsive to changes in the views of their constituents. First, the re-election incentives for legislators to respond to changes in public opinion actually appear to be surprisingly modest, possibly because many voters have only a dim awareness of legislators' roll call positions. Indeed, Canes-Wrone, Brady, and Cogan (2002) finds that a one standard deviation moderation in legislators' roll call behavior only leads to a 2% increase in legislators' vote shares. Second, voters appear to punish candidates that reposition their views on individual issues (Van Houweling and Tomz, 2011). This could reduce candidates' incentives to respond to

changes in constituent opinion on individual issues.

There is a clear consensus among scholars that more liberal states are more likely to elect senators that take liberal roll call positions (Erikson and Wright, 2000; Ansolabehere, Snyder, and Stewart, 2001; Clinton, 2006). However, there is no consensus about whether senators respond to changes in the public's policy preferences in their state. Some studies find little or no evidence of dynamic responsiveness. For instance, Poole (2007, 435) finds that legislators do not change their positions over the course of their careers and thus "die in their ideological boots." Along these lines, Krimmel, Lax, and Phillips (2015) finds no evidence that senators change their positions on gay rights in response to changes in their constituents views.

In contrast, several other studies find evidence of modest dynamic responsiveness in the Senate. Miler (Forthcoming) finds that newly elected senators that had previously served in the House shift their positions toward the statewide median voter. Also, Buttice and Highton (Forthcoming) finds that senators shift their positions in response to changes in presidential voting patterns in their state. However, the implications of this study are unclear because presidential vote shares may not clearly reflect public opinion in earlier eras, especially in the one party south (Levendusky, Pope, and Jackman, 2008; Caughey and Warshaw, 2015). So it's unclear whether Buttice and Highton (Forthcoming) are observing responsiveness to changes in public opinion, partisanship, voting behavior, or some other latent quantity.

Changes in Responsiveness Over the Electoral Cycle

There are strong theoretical reasons to believe that Senators' political incentives change over the course of the electoral cycle. Most importantly, a variety of observational studies have found that voters overweight recent events when they evaluate incumbents' performance (e.g., Achen and Bartels, 2004). Several experimental studies have also found strong evidence that voters overweight recent economic performance (Huber, Hill, and Lenz, 2012; Healy and Lenz, 2014). For instance, Healy and Lenz (2014) find that voters focus on the election-year economy because it is more accessible than information about the economy across the entire electoral cycle. They suggest that this "reflects a general tendency for people to simplify retrospective assessments by substituting conditions at the end for the whole."

This tendency to overweight recent events is likely to be even stronger for assessments of legislators' roll call voting behavior than for the economy due to voters' general lack of information on legislators' roll call positions. Franklin (1993) finds that respondents to the 1988-1992 Senate Election Study were more likely to be able to place legislators on an ideological scale in the last two years of their term.

Voters' tendency to overweight recent events may give legislators freedom to take unpopular positions early in their terms. In addition, it is likely to incentivize politicians to consider the views of their constituents more as elections grow more proximate. Indeed, a variety of qualitative evidence indicates that Senators have a general heuristic that they should focus more on public opinion in the last two years of their term. For example, one 18-year Senate veteran told Richard Fenno: "My life [as a Senator] has a six-year cycle to it... I say in the Senate that I spend four years as a statesman and two years as a politician.

You should get cracking as soon as the last two years open up. You should take a poll on the issues, identify people to run my campaign in different parts of the state, raise money, start my PR, and so forth" (Fenno, 1982, 29).

While there is little existing evidence about how responsiveness changes over the electoral cycle, previous studies have found that Senators moderate their roll call behavior as elections approach, presumably because they are shifting their behavior toward the median voter in their state (Albouy, 2011; Amacher and Boyes, 1978; Elling, 1982; Lindstädt and Vander Wielen, 2011; Thomas, 1985). These studies show that Democrats tend to shift to the right in the last two years of their term, while Republicans tend to shift to the left.

Research Design

I examine dynamic responsiveness in the United States Senate using new estimates of citizen policy conservatism for each state and year that are derived from the survey responses of approximately 730,000 individuals to over 150 domestic policy questions fielded between 1950 and 2012 (Caughey and Warshaw, 2015). I also estimate a dynamic ideal point model to estimate how Senators' ideal points change over time (Martin and Quinn, 2002). With these new data in hand, I can test whether Senators are responsive to changes in the views of their constituents using a time-series cross-sectional (TSCS) approach, which enables me to exploit within-legislator variation in public opinion and roll call behavior. Moreover, I can examine how Senators' roll call behavior changes over the course of the electoral cycle.

Citizen Policy Conservatism: 1950-2012

I follow Erikson, Wright, and McIver (1993), Stimson, MacKuen, and Erikson (1995), and many others, in summarizing public opinion on a single dimension, which I label *citizen economic policy conservatism* to distinguish it from ideological self-identification.[4] To estimate citizen policy conservatism in each state-year, I apply the dynamic, hierarchical group-level IRT model developed by Caughey and Warshaw (2015, 2016), which estimates the average policy conservatism of the population in each state. This approach builds upon three important approaches to modeling public opinion: item-response theory, multilevel regression and poststratification, and dynamic measurement models (See Appendix A for details of the measurement model).

My public opinion data consists of survey responses to over 150 domestic economic policy questions fielded between 1950 and 2012. The questions cover traditional economic issues such as taxes, social welfare, and labor regulation. The responses of over 700,000 different Americans are represented in the data. I model opinion in groups defined by states and a set of demographic categories (e.g., race and gender). In order to mitigate sampling error for small states, I model the state effects in the first time period as a function of state *Proportion Evangelical/Mormon.* The inclusion of state attributes in the model partially pools information across similar geographical units in the first time period, improving the efficiency of state estimates (e.g., Park, Gelman, and Bafumi, 2004, 2006). I drop *Proportion Evangelical/Mormon* after the first period because I found that the state intercept in the

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[4] Relative to conservatism, liberalism involves greater government regulation and welfare provision to promote equality and protect collective goods, and less government effort to uphold traditional morality and social order at the expense of personal autonomy. Conversely, conservatism places greater emphasis on economic freedom and cultural traditionalism (Ellis and Stimson, 2012, 3–6).

previous period tends to be much more predictive than state attributes. To generate annual estimates of average opinion in each state, I weighted the group estimates to match the groups' proportions in the state population, based on data from the U.S. Census (Ruggles et al., 2010).[5]

Dynamic Ideal Points

In order to examine how Senators' ideal points vary over the course of the electoral cycle, I need a dynamic estimate of their ideal points that is not assumed to either stay constant over-time or evolve via a smooth time trend. Therefore, both Common Space and DW-Nominate scores are inappropriate for this application (Caughey and Schickler, 2014; Buttice and Highton, Forthcoming). Instead, I measure legislators' roll call behavior based on two different dynamic ideal point models. First, I use Nokken-Poole ideal points (Nokken and Poole, 2004).[6] Second, I use dynamic Bayesian ideal points (Martin and Quinn, 2002). I orient both sets of ideal points so that positive values indicate more conservative roll call voting behavior.

The Nokken-Poole ideal points are similar to DW-Nominate scores. However, they allow the maximum amount of movement of Senators from Congress to Congress (Nokken and Poole, 2004). The model proceeds in two steps. First, it runs a two-dimensional DW-NOMINATE model. This model assumes that every legislator has the same ideal point

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[5] The most conservative states are in the Great Plains, while New York, California, and Massachusetts are always among the most liberal states. While most states have remained generally stable in their relative conservatism (Erikson, Wright, and McIver, 2006, 2007), a few states' policy conservatism has shifted substantially over time. Southern states such as Mississippi and Alabama have become more conservative over time, while states in New England have become more liberal.

[6] Nokken and Poole (2004) used this method to examine how much party switchers changed their ideal points between Congresses.

throughout his or her career. Second, it estimates an ideal point for every legislator in every Congress using the item parameters from the DW-Nominate model. This approaches enables analysts to evaluate changes in the ideal points of Senators against the background of the fixed cutting lines. This approach yields estimates of Senators ideal points that can change inter-temporally, but are comparable over time.

A weakness of the Nokken-Poole scores is that they are only available for each Congress, rather than for each year. This makes it difficult to precisely match them to my annual estimates of citizen policy conservatism. In order to obtain annual estimates of Senators' ideal points, I estimate dynamic ideal points for each Senator between between 1950-2012 using a Bayesian model that allows each Senator's ideal point to follow a random walk over time (Martin and Quinn, 2002; Clinton, Jackman, and Rivers, 2004).[7] I use the Policy Agendas Project to identify roll calls on economic issues in order to ensure that I am capturing the first dimension of roll call votes.[8] It is important to note that a weakness of this dynamic ideal point model is that it shrinks the inter-temporal changes in legislators' ideal points, which could bias the estimated "treatment effect" of election proximity toward zero (Lindstädt and Vander Wielen, 2011). Thus, the "true" treatment effects of public opinion and election proximity are likely to be larger than we observe using these dynamic ideal points.

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[7] In this approach, changes in Senators' ideal points between congresses are assumed to follow a normal distribution centered at zero. Legislators may move to the left or right, but their expected ideal point in a given congress is their location in the previous congress. A Senator's ideal point in congress t is a weighted combination of their ideal point in the previous congress and the ideal point implied by their voting record in congress t (Caughey and Schickler, 2014). The user-imposed prior regarding the variance in Senators' ideal points in congress t around their location in the previous time period identifies the model inter-temporally.

[8] I estimate Bayesian dynamic ideal points for all Senators between 1950-2012 with MCMCpack in R. I estimate the model based on a random sample of 150 economic roll call votes in each year. This introduces a bit more variance into my estimates, but it substantially reduces the model's running time.

Measuring Dynamic Responsiveness

My primary research design examines how legislators' roll call behavior evolves in response to changes in public opinion in their state. The advantage of my time-series cross-sectional (TSCS) dataset is that it enables me to exploit within-legislator variation in opinion and ideal points, and thus control for the persistent, time invariant characteristics of each legislator. I use an unbalanced panel model with fixed effects (FEs) for Senator and session/year x party interactions, which controls for time-invariant characteristics of each legislator and for time-specific effects common to all legislators in each party (Angrist and Pischke, 2009).[9]

Election Proximity

I measure the proximity of elections using a dummy variable indicating whether a Senator was in the last two years of his/her term. Due to endogeneity concerns, I do not differentiate between Senators that actually ran for re-election and those that retired. Indeed, many Senators do not make the decision about whether to run for re-election until late in their term, and the decision to run for re-election could, of course, be influenced by public opinion. The results are substantively similar if I use a continuous measure of the number of years until a Senator's next election.

Dynamic Responsiveness

In this section, I present the main results of my analysis of the effect of citizen policy conservatism on Senators' ideal points, and how this effect varies over the course of the

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[9] I obtain substantively similar results using a lagged dependent variables (LDVs) to capture unobserved omitted variables in each unit (De Boef and Keele, 2008; Beck and Katz, 2011).

electoral cycle. I present results using both Nokken-Poole and Bayesian ideal points. Both sets of ideal points are oriented so that positive vales represent more conservative roll call behavior. In addition, both sets of ideal points are standardized to yield more straightforward and comparable substantive interpretations. Finally, I lag the measures of public opinion by one time period to avoid concerns of reverse causation whereby constituents are changing their ideological positions based on how their senators vote (Lenz, 2013).

The first two columns of Table 1 show the results of a series of models using Nokken-Poole ideal points as the dependent variable. The results in column (1) indicate that a one standard deviation shift to the left in citizen policy conservatism causes Senators to shift about .024 standard deviations to the right. The results in column (2) indicate that legislators are roughly twice as responsive to public opinion in the last two years of their term. Indeed, in the last two years of their term, a one standard deviation shift to the right in citizen policy conservatism causes Senators to shift about .036 standard deviations to the right.

The next two columns show the results of a series of models with Bayesian dynamic ideal points as the dependent variable (Martin and Quinn, 2002; Clinton, Jackman, and Rivers, 2004). Column (3) indicates that a one standard deviation shift to the right in citizen policy conservatism causes Senators to shift about .030 standard deviations to the right. Similarly to the results using Nokken-Poole ideal points, Column (4) indicates that Senators' dynamic ideal points are about twice as responsive to public opinion in the last two years of their terms as in the first four years.

Overall, this analysis demonstrates two new facts about Senators' roll call behavior. First, Senators are modestly responsively to public opinion. While the magnitude of these shifts is

	Dependent variable:			
	Nokken-Poole Ideal Points		Dynamic Ideal Points	
	(1)	(2)	(3)	(4)
Citizen Policy Conservatism$_{t-1}$	0.024**	0.016	0.030***	0.022***
	(0.012)	(0.012)	(0.008)	(0.008)
Policy Cons.$_{t-1}$ x Last 2 years		0.020**		0.020***
		(0.008)		(0.006)
Last two years of term	−0.086***	−0.088***	−0.059***	−0.063***
	(0.012)	(0.012)	(0.008)	(0.008)
Last 2 years x Dem.	0.127***	0.135***	0.087***	0.096***
	(0.016)	(0.016)	(0.011)	(0.011)
Constant	0.140**	0.142**	0.252***	0.254***
	(0.069)	(0.069)	(0.050)	(0.050)
Fixed Effects for Session/Year	Y	Y	Y	Y
Fixed Effecs for Senator	Y	Y	Y	Y
Observations	3,103	3,103	6,136	6,136
R^2	0.968	0.968	0.969	0.969
Adjusted R^2	0.961	0.961	0.965	0.966

Note: *p<0.1; **p<0.05; ***p<0.01

small, to my knowledge, this is the first study that has shown that Senators are responsive to changes in public opinion. Second, legislators are much more responsive to public opinion in the last two years of their terms than in the first four years.

Responsiveness to Issue-Specific Public Opinion

An important advantage of the analysis in the previous section is that it links changes in public opinion to changes in legislators' roll call behavior (ideal points). This approach enables

me to use time-series cross-sectional models to account for unobserved omitted variables that might weaken causal inferences. However, a disadvantage is that it makes potentially strong assumptions about the dimensionality of both public opinion and legislators' roll call behavior (Broockman, 2016). In addition, it is impossible to evaluate the spatial proximity between citizen policy conservatism and legislators' ideal points since they are measured on different scales (Lewis and Tausanovitch, 2013; Jessee, 2016).

An alternative approach is to examine the association between public opinion on individual issues and legislators' roll call votes on those issues. While the bulk of the literature on representation focuses on the link between an aggregate measure of citizen policy conservatism and legislations' one dimensional ideal points, there is also a vibrant literature that argues that representation should be examined on an issue-by-issue basis (see Broockman, 2016, for an overview).[10] An advantage of this approach is that it makes fewer assumptions about the dimensionality of public opinion or legislators' roll call behavior. It also makes it possible to compare public opinion and roll call votes on the same scale (Lax and Phillips, 2009a; Matsusaka, 2010). In the remainder of this section, I examine the robustness of my results by evaluating how election proximity affects the association between public opinion on over 72 individual issues with senators' roll call votes on those issues.

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[10] A variety of studies have examined representation on an issue-by-issue basis. Gilens (2005) uses an issue-by-issue approach to examine variation in policy representation across income groups. In the state politics literature, Lax and Phillips (2009a, 2011), and Matsusaka (2010) examine the congruence between state policies and issue-specific public opinion. In the literature on dyadic representation, Kastellec, Lax, and Phillips (2010) examine the link between public opinion and Senators' votes on Supreme Court nominations, while Krimmel, Lax, and Phillips (2015) examines the link between issue-specific public opinion and legislators' votes on gay rights roll calls.

Measuring Issue-Specific Public Opinion

In order to examine representation on an issue-by-issue basis, the first task is to measure public opinion on a variety of issues that have been voted upon in Congress. In order to do this, I use a technique called multi-level regression with post-stratification (MRP) to estimate public opinion in every state on a wide range of issues considered by the Senate in the 97th-111th Congresses. This approach employs Bayesian statistics and multilevel modeling to incorporate information about respondents' demographics and geography in order to generate accurate state-level estimates of public opinion (Park, Gelman, and Bafumi, 2004; Lax and Phillips, 2009b).

There are two stages to the MRP model (see Appendix B for a full description of the model). In the first stage, I estimate each individual's preferences as a function of his or her demographics and state. This approach allows individual-level demographic factors and geography to contribute to my understanding of state opinion. In the second stage, I use the multi-level regression to make a prediction of public opinion in each demographic-geographic sub-type. The estimates for each respondent demographic geographic type are then weighted by the percentages of each type in the actual state populations. Finally, these predictions are summed to produce an estimate of public opinion in each state.

The universe of issues in my analysis are roll call votes in the 97-111th Congresses (1981-2010) that were included in the Americans for Democratic Action (ADA)'s annual scorecard, listed as "key votes" by Congressional Quarterly, and all Supreme Court nominations during this period.[11] To build my public opinion estimates, I obtained all available public opinion data on these roll call votes. I used several rules to govern my collection of survey data.

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[11] I also only used roll calls that did not occur during lame duck congressional sessions.

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First, I only gathered data from surveys where the respondents could be matched to states. Second, I only used data from surveys where the question is specific enough to be matched to an individual roll call vote. For instance, I did not use general data on support for abortion rights to assess whether legislators are responsive to their constituents' preferences on a partial birth (late-term) abortion vote. Instead, I only used survey data about constituent preferences on partial birth abortions. Third, I only used surveys that were collected within a two-year period before the roll call vote to ensure that public opinion is measured prior to my outcome variable. This also minimizes concerns of reverse causation (Lenz, 2013). Fourth, I did not use exit polls or other surveys that are limited to just voters. Fifth, I only estimated public opinion for issues with survey data available for at least 2,000 citizens.[12]

I was able to obtain public opinion data for 72 roll call votes from the 97th-111th Congresses (Figure 1).[13] These roll calls are not a random sample of the congressional agenda and so some caution must be taken in generalizing the findings. However, the bills were not purposefully selected on substantive grounds or by the degree to which they line up with ideology or opinion measures. If an issue was voted on in multiple Congresses, each roll call was included separately. For instance, there are votes from both the 108th and 109th Congresses on a gay marriage amendment. The roll calls come from a variety of issue areas, including economic, social, and foreign policy issues, as well as a number of Supreme Court nominations. Some of the roll calls sought to move the status quo in a conservative direction,

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[12] This somewhat conservative cut-off is based on the validation analysis in Lax and Phillips (2009b), which indicates that national samples larger than 1,500 citizens are sufficient to estimate public opinion in each state with a relatively high degree of accuracy.

[13] I used survey data from the following organizations: the 2000 and 2004 Annenberg National Election Surveys, the 2006, 2007, 2008, 2009, and 2010 Cooperative Congressional Election Surveys, the 2000, 2002, and 2009 American National Election Surveys (I only used ANES surveys that employed random-digit dial sampling rather than cluster-sampling.), various CBS News/New York Times polls, ABC News/Washington Post polls, Gallup polls, Pew Foundation polls, NBC News polls, and LA Times polls.

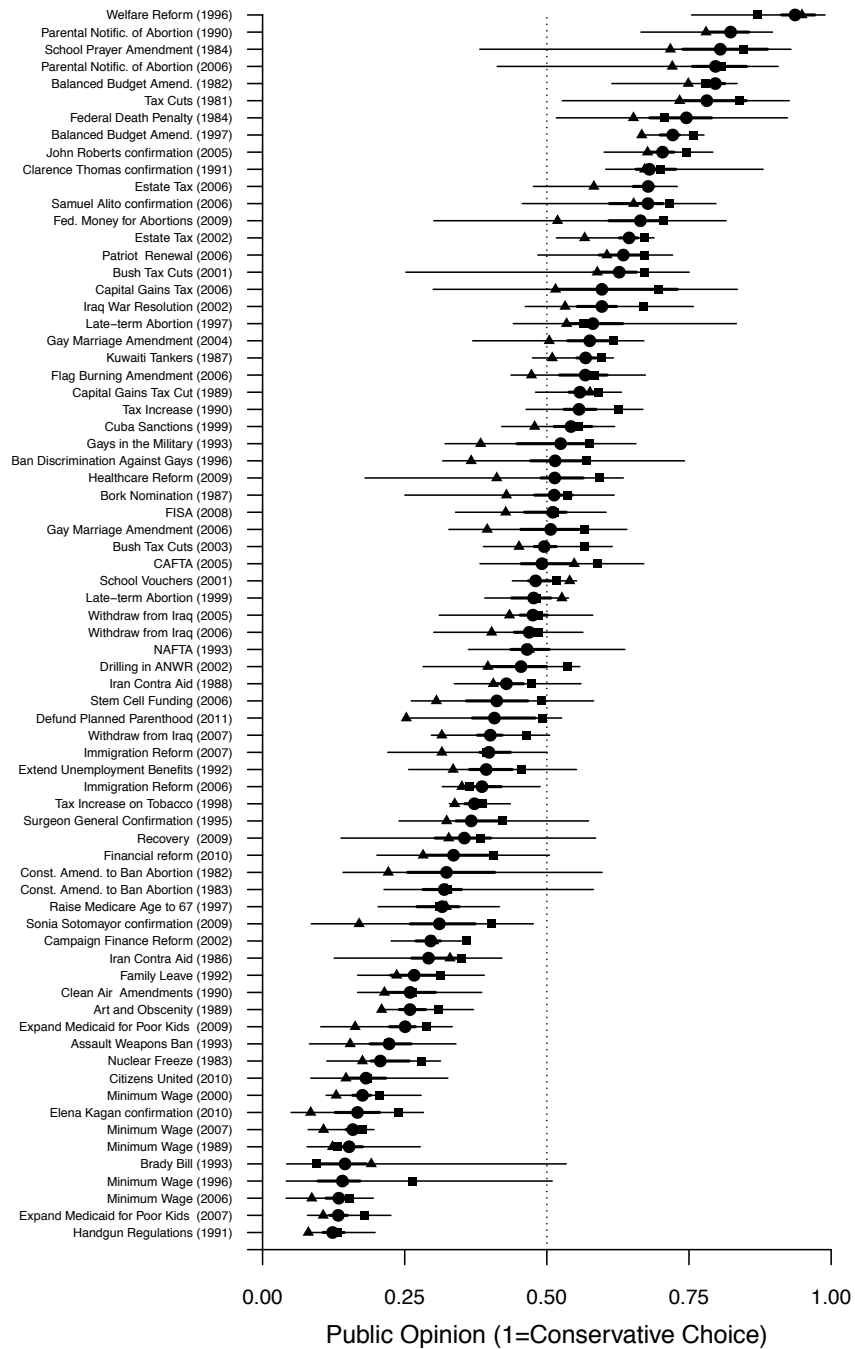while others sought to move it in a liberal direction.

Figure 1: Public Opinion on each issue - This figure shows MRP estimates of public opinion. The dots show opinion in the median state, the thick lines show opinion in the middle 50 percent of states, and the thin lines shows opinion over the entire range. To illustrate how opinion varies within states across issues, the triangles show opinion in New York and the squares show opinion in Texas.

Figure 1 shows public opinion for each of the roll calls in my dataset. It indicates that opinion varies significantly across issues. On some issues, there are clear liberal majorities, while there are strong conservative majorities on other issues. The most liberal opinion majorities are on handgun registration and expanding Medicaid for poor children. On the other end of the spectrum, the most conservative opinion majorities are for welfare reform, parental notification of abortion, and school prayer.

Issue-Specific Results

I run two sets of multi-level logistic regression models to examine the impact of election proximity on Senators' roll call votes. First, I analyze the association between public opinion and roll call votes ("responsiveness") through a set of models that examine whether variation in public opinion is correlated with roll call votes after controlling for other factors (Achen, 1978; Egan, 2005). In these models, I interact the proximity of the next election with public opinion to see how the association between public opinion and roll call votes varies over the electoral cycle. I also include legislators' party ID, as well as issue and legislator random/fixed effects.

It could be the case that members of Congress are responsive to public opinion, but systemically more liberal, or conservative, than their constituencies (Achen, 1978; Erikson and Wright, 2000). Thus, the concept of *congruence* assesses whether legislators follow the preferences of a majority of their constituents on individual roll call votes (Egan, 2005; Krimmel, Lax, and Phillips, 2015; Weissberg, 1979). I analyze congruence through a set of models that examine whether election proximity affects the probably that legislators' votes

are congruent with the views of the majority of citizens in their state.

	Dependent variable:				
	Rollcall Vote				
	(1)	(2)	(3)	(4)	(5)
Issue-Specific Public Opinion	5.01***	4.69***	4.80***	4.84***	2.96***
	(0.0004)	(0.62)	(0.62)	(0.62)	(0.90)
Public Opinion * Last two years	1.85***	1.78***	1.45***	1.39***	1.68***
	(0.0004)	(0.45)	(0.47)	(0.47)	(0.51)
Last two years	−0.76***	−0.73***	−0.79***	−0.79***	−0.86***
	(0.0004)	(0.23)	(0.23)	(0.23)	(0.25)
Last two years:Democrat			0.42**	0.44**	0.39*
			(0.20)	(0.19)	(0.21)
Democrat		−5.15***	−5.30***	−5.29***	
		(0.38)	(0.39)	(0.38)	
Democratic Presidential Share		−7.06***	−7.02***	−7.23***	
		(1.14)	(1.14)	(1.13)	
Constant	−2.84***	3.85***	3.85***	3.89***	0.18
	(0.0004)	(0.67)	(0.67)	(0.66)	(1.42)
Random Effects for Party x Issue	X	X	X	X	
Random Effects for Legislator	X	X	X	X	
Fixed Effects for Party x Issue					X
Fixed Effects for Legislator					X
Observations	7,094	7,094	7,094	7,275	7,094
Log Likelihood	−2,285.02	−2,208.13	−2,205.86	−2,252.60	−1,617.77
Akaike Inf. Crit.	4,586.04	4,436.27	4,433.73	4,527.21	4,083.54
Bayesian Inf. Crit.	4,640.97	4,504.94	4,509.26	4,603.02	

Note: *p<0.1; **p<0.05; ***p<0.01

Table 2: Senators' Responsiveness to Public Opinion

Is there a stronger association between issue-specific public opinion and roll call votes

when elections are proximate? Using a model with random effects for legislators and bills,
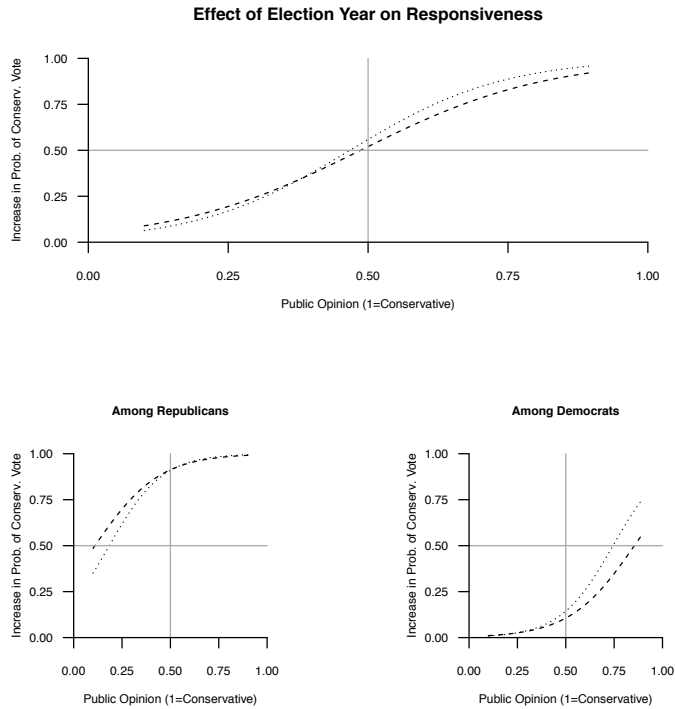
Figure 2: Issue-Specific Responsiveness - This figure shows the association between variation in issue-specific public opinion and Senators' roll call votes. The dotted lines show the association between opinion and votes during the last two years of senators' terms and the dashed lines show the association during the remainder of the term.

Columns 1-4 of Table 2 show that, on average, Senators are more responsive to public opinion in the last two years of their term. Indeed, Senators are about 25 percent more responsive when elections are proximate. These results are robust across modeling specifications. In addition, Column 5 shows that a model with fixed effects for legislators and bills yields similar results. Figure 2 plots the substantive meaning of these results. When 75 percent of the public supports a liberal policy, there is a 38 percent probability that Republicans in the last two years of their term will support the liberal policy, compared with just a 30 percent probability of supporting the liberal policy for Republicans that are not in the last two years

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of their term. Likewise, when 75 percent of the public supports a conservative policy, there is a 51 percent probability that Democrats in the last two years of their term will support the conservative policy, compared with a 35 percent probability of supporting the conservative policy for Democrats that are not in the last two years of their term.

Table 3:

	Dependent variable:	
	Congruent	
	(1)	(2)
Size of Issue-Specific Majority	5.90***	7.29***
	(0.58)	(0.71)
Last two years	0.13**	0.10
	(0.07)	(0.07)
Democrat	0.73	
	(0.46)	
Constant	−3.46***	−17.50
	(0.47)	(6,522.65)
Random Effects for Party x Issue	X	
Random Effects for Legislator	X	
Fixed Effects for Party x Issue		X
Fixed Effects for Legislator		X
Observations	7,094	7,094
Log Likelihood	−3,267.38	−2,802.42
Akaike Inf. Crit.	6,550.76	6,448.85
Bayesian Inf. Crit.	6,605.70	
Note:	*p<0.1; **p<0.05; ***p<0.01	

Table 4: Congruence with Public Opinion, by Party

Moving beyond responsiveness, are Senators' votes congruent with the preferences of a majority of their constituents (Table 4)?[14] Columns 1 and 2 show the average effect of election

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[14] On the whole, about 60 percent of the roll call votes in my dataset are congruent with the views of

proximity on congruence across all Senators. Column 1 uses random effects and column 2 uses fixed effects. Both models indicate that Senators are somewhat more congruent with issue-specific public opinion in the last two years of their term. This is broadly consistent with evidence from a variety of previous studies that Senators tend to moderate their behavior in the last two years of their term, since more moderate positions are generally more proximate to the location of the median voter.

Conclusion

The foundation of representative democracy is the assumption that citizens' preferences should correspond with, and inform, elected officials' behavior. Thus, in order to understand how well our democracy is functioning, it is crucial to know the degree to which legislators actually follow the views of their constituents. In this study, I show that there is a significant relationship between public opinion and Senators' roll call votes, both at an aggregate level and on individual issues. But my results also indicate that the aggregate effect of public opinion is still small.

In addition, I resolve earlier ambiguities in the literature about the relationship between election proximity and responsiveness over the past six decades. Three decades after Fenno's pathbreaking study of representation in the Senate, the literature on the relationship between election proximity and responsiveness has remained ambiguous. This study demonstrates

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issue-specific opinion majorities. The level of congruence between legislators' roll calls and the majority of the public varies significantly across issues though. To put the level of congruence I find in context, it can be compared to studies of specific issue areas in Congress. Krimmel, Lax, and Phillips (2015) find that 64 percent of recent Congressional roll calls on gay rights are congruent with majority opinion. Kastellec, Lax, and Phillips (2010) find that Senators' Supreme Court votes are congruent with public opinion 79 percent of the time.

that over the past six decades responsiveness in the United States Senate has predictably varied over the course of the electoral cycle. My results provide strong support for the notion that Senators are more attentive to their constituents' views later in the electoral cycle. The results here complement earlier findings that Senators move toward the electoral middle when elections approach (Albouy, 2011; Amacher and Boyes, 1978; Elling, 1982; Lindstädt and Vander Wielen, 2011; Thomas, 1985), as well as more recent findings that appropriations are influenced by the electoral cycle (Shepsle et al., 2009).

Future work should examine other areas such as legislative effort and constituent service to examine whether they also vary over the electoral cycle. It should also examine other legislative contexts such as state legislatures, where many states vary term lengths across upper house districts (Titiunik, Forthcoming).[15]

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[15] Of course, it is possible that the lower salience of elections at the state-level reduces the incentives for legislators there to respond to changes in public opinion.

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Appendix A: Methodology for Estimating Citizen Policy Conservatism

The lack of a time-varying measure of citizens' policy preferences in each state has been one of the main barriers to the study of representation in Congress. To overcome this challenge, I apply the dynamic hierarchical group-level IRT model developed by Caughey and Warshaw (2015, 2016), which estimates the average policy conservatism of defined subpopulations (e.g., non-urban whites, urban whites, and blacks in each state). This model adopts the general framework of item-response theory (IRT). In an IRT model, respondents' question responses are jointly determined by their score on some unobserved trait—such as their economic policy conservatism—and by the characteristics of the particular question. The relationship between responses to question q and the unobserved trait θ_i is governed by the question's threshold κ_{qt}, which captures the base level of support for the question, and its dispersion σ_q, which represents question-specific measurement error. Under this model, respondent i's probability of selecting the conservative response to question q is

$$\pi_{iq} = \Phi\left(\frac{\theta_i - \kappa_{qt}}{\sigma_q}\right),\tag{1}$$

where the normal CDF Φ maps $(\theta_i - \kappa_{qt})/\sigma_q$ to the $(0,1)$ interval.[1] The model assumes that greater conservatism (i.e., higher values of θ_i) increases respondents' probability of answering in a conservative direction. The strength of this relationship is inversely proportional to σ_q, and the threshold for a conservative response is governed by κ_{qt}. The probability that a

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[1] A common alternative way of writing the model in Equation (1) is $\Pr(y_{iq} = 1) = \Phi(\beta_q \theta_i - \alpha_q)$, where $\beta_q = 1/\sigma_q$ and $\alpha_q = \kappa_{qt} \times \beta_q$.

randomly sampled member of group g correctly answers item q is

$$\pi_{gq} = \Phi\left(\frac{\bar{\theta}_g - \kappa_{qt}}{\sqrt{\sigma_q^2 + \sigma_\theta^2}}\right), \tag{2}$$

where σ_θ is the standard deviation of θ_i within groups. Equation (2) is connected to the data through the sampling model

$$s_{gq} \sim \text{Binomial}(n_{gq}, \ \pi_{gq}), \tag{3}$$

where n_{gq} is group g's total number of non-missing responses to question q and s_{gq} is the number of those responses that are conservative.[2] The estimates of $\bar{\theta}_g$ can be poststratified into estimates of the average economic conservatism in each state (cf. Park, Gelman, and Bafumi (2004)).

To address sparseness in the matrix of survey data, I use a dynamic linear model to smooth the estimated group means across both time and states.

$$\bar{\theta}_{gt} \sim N(\delta_t \bar{\theta}_{g,t-1} + \xi_t + \mathbf{x}'_{g\cdot}\gamma_t, \ \sigma_{\bar{\theta}t}^2), \tag{4}$$

where $\bar{\theta}_{g,t-1}$ is g's mean in the previous year, ξ_t is a year-specific intercept, and $\mathbf{x}_{g\cdot}$ is a vector of attributes of g (e.g., its state or party). Each group-year mean is thus modeled as a function of the group's mean in the previous year, year-specific changes common to all groups, and changes in relative conservatism of groups with similar characteristics (i.e., the

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[2] Following Ghitza and Gelman (2013) and Caughey and Warshaw (2015, 202–3), I adjust the raw values of s_{gq} and n_{gq} to account for survey weights and for respondents who answer multiple questions.

same party or state). The posterior estimates of $\bar{\theta}_{gt}$ are a thus compromise between this prior and the likelihood implied by Equations (2) and (3), with the relative weight placed on the likelihood determined by the prior standard deviation $\sigma_{\bar{\theta}t}$, which is estimated from the data and allowed to evolve across years. When a lot of survey data are available for a given year, the likelihood will dominate. If no survey data are available at all, the prior acts as a predictive model that imputes $\bar{\theta}_{gt}$.

The dynamic group-level IRT model estimates opinion in groups defined by states and demographic groups. To generate annual estimates of average opinion in each state, the survey data are pre-weighted to match raked targets for gender and education level in each state public, based on data from the U.S. Census (Ruggles et al., 2010). This model produces estimates of the ideology of non-urban whites, urban whites, and blacks in each state. Finally, the estimates are aggregated up to the national level based on post-stratification weights from the PUMS data available from the Census.

Appendix B: MRP Methodology for Estimating Issue-Specific Public Opinion

Multilevel regression and poststratification (MRP) models employ Bayesian statistics and multilevel modeling to incorporate information about respondents' demographics and geography in order to estimate public opinion in each geographic subunit (Park, Gelman, and Bafumi, 2004; Lax and Phillips, 2009; Warshaw and Rodden, 2012). Specifically, each individual's survey responses are modeled as a function of demographic and geographic predictors, partially pooling respondents across state to an extent determined by the data (see Gelman and Hill, 2007; Jackman, 2009 for more about multilevel modeling). The state-level effects are modeled using additional state-, and region-level predictors, such as states' median-income level and religiosity. Thus, all individuals in the survey yield information about demographic and geographic patterns, which can be applied to all state estimates. The final step is poststratification, in which the estimates for each demographic-geographic respondent type are weighted (poststratified) by the percentage of each type in the actual state population.

There are two stages to the MRP model. In the first stage, I estimate each individual's preferences as a function of his or her demographics and and state (for individual i, with indexes r, e, a, p, s, and z for race, gender, education category, poll, state, and region, respectively). This approach allows individual-level demographic factors and geography to contribute to my understanding of state opinion. Moreover, the model incorporates both within and between state geographic variation. I facilitate greater pooling across states by including in the model several state-level variables that are plausibly correlated with public

opinion. For example, I include the percentage of people in each state that are evangelicals or Mormons.

I incorporate this information with the following hierarchical model for respondent's responses:

$$Pr(y_i = 1) = logit^{-1}(\gamma_0 + \alpha_{r[i]}^{race} + \alpha_{g[i]}^{gender} + \alpha_{e[i]}^{edu} + \alpha_{a[i]}^{age} + \alpha_{s[i]}^{state} + \alpha_{p[i]}^{poll})$$

$$\text{where:}$$
$$\alpha_{r[i]}^{race} \text{ for r} = 1, 2$$
$$\alpha_{g[i]}^{gender} \text{ for g} = 1, 2 \tag{5}$$
$$\alpha_{e[i]}^{edu} \text{ for e} = 1, \ldots, 5$$
$$\alpha_{p[i]}^{poll} \text{ for p} = \text{the number of polls for each issue}$$

That is, each individual-level variable is modeled as drawn from a normal distribution with mean zero and some estimated variance. Following previous work using MRP, I assume that the effect of demographic factors does not vary geographically. I allow geography to enter into the model by adding a state level to the model, and giving each state a separate intercept.

The state effects are modeled as a function of the region into which the state falls, the percentage of the state's residents that are union members, the state's percentage of evangelical or Mormon residents, the state's average income, and the percent of the state's residents that live in urban areas.

$$\alpha_s^{state} \sim N(\alpha_z^{region} + \beta_1 * union_s + \beta_2 * religion_s + \beta_3 * income_s + \beta_4 * urban_s, \sigma_s^2)$$

$$\tag{6}$$

$$\text{for s} = (1, \ldots, 51)$$

The region variable is, in turn, another modeled effect. I group states into regions based on

their general ideology and vote in presidential elections.

$$\alpha_{z[i]}^{region} \sim N(0, \sigma_z^2)$$

$$\text{for } z = (1, \ldots, 6) \tag{7}$$

I estimate the model using the GLMER function in R (Bates 2005).

Poststratification

For any set of individual demographic and geographic values, cell c, the results above allow me to make a prediction of public opinion. Specifically, c is a function of the relevant predictors and their estimated coefficients. Next, I weight these estimates by the percentages of each type in the actual state populations. I calculate the necessary population frequencies using PUMS "5-Percent Public Use Microdata Sample" from the census, which has demographic information for 5 percent of each state's voting- age population (Park, Gelman, and Bafumi, 2004; Lax and Phillips, 2009).

Validation

The MRP estimates of state-level public opinion are generally highly correlated with the raw "disaggregated" state-level estimates of public opinion from the various surveys. Across all issues, the MRP estimates of public opinion are correlated with the disaggregated estimates at 0.91. For roll calls with smaller sample sizes, however, there are sometimes smaller correlations between the disaggregated and MRP estimates. The advantage of MRP is that it enables me to accurately estimate public opinion for issues, such as closing Guantanamo,

or geographic areas, such as Montana and North Dakota, where survey data is sparse.

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