Climate voting in Congress: The power of public concern

Abstract

In this study, we test the relationship between congressional votes and public concern about climate change. In the US, very few constituents know and understand climate policy, prioritize it as a political topic, or let their voting decisions depend on it. In these conditions, we may not expect representatives to take public concern about climate change into account in their voting decisions. Still, even after controlling for the presence of interest groups, campaign finance, and legislators' party affiliation and ideology, we find a consistent link between public opinion and votes on cap-and-trade legislation in the House (and to a lesser degree in the Senate). The same is true when we simulate public opinion based on pre-vote district characteristics. This finding raises questions about the nature of public concern on climate change, and representation in Congress in general.

Keywords: climate change, public opinion, Congress, representation, constituency, responsiveness

1 Introduction

It is a barely controversial observation that federal level climate policy in the US has so far been much less far-reaching than that in other industrialized countries, especially European ones. Several studies have been dedicated to the search for potential causes of this relative inaction, and weak public concern about climate change has regularly been cited as a prime suspect. For example, Steurer (2003) has argued that President George W. Bush was able to retreat from the Kyoto Protocol in part because of public disinterest in the climate issue. Indeed, Harrison & Sundstrom (2010) find it telling that this decision incited "larger protests across Europe than in the United States itself". Still, some studies of US climate policymaking do not cite any influence of public opinion at all. They ascribe the lack of ambitious federal climate policy to pressure from industrial interest groups, or to weak environmental organizations (Bryner 2008, Skodvin & Andresen 2009).

In other words, there does not seem to be solid agreement among authors as to whether and how much public opinion has contributed to the relative "climate conservatism" of the US federal government.

That includes the effect of public opinion on congressional decision-making, which is the focus of this study. As we will see, this lack of agreement also reflects a large degree of uncertainty in the broader literature on political responsiveness. As a result, studying climate politics could also help fill some of the gaps in our current understanding of congressional representation in general.

1.1 Political Representation

The idea that legislators' decisions are primarily based on constituents' preferences is a key concept in representation theory. According to what Gilens & Page (2014) call theorists of "majoritarian electoral democracy" (e.g. Dahl 1956, Downs 1957), elected politicians primarily follow the preferences of their constituents when making policy decisions. Theory points to two main pathways through which public opinion may be reflected in the votes of their representatives: a process of selection and a process of influence. First, constituents can elect those representatives they think will best represent their interests. Second, voters have at their disposal a range of levers to change the voting behavior of their representatives after the elections. For the general public, the most important of these levers is to threaten with voting for another candidate in the next election (Canes-Wrone et al. 2002).

According to Page (1994), the research on representation of public opinion "now encompasses hundreds of articles, as well as major books." Most studies that have tested the relationship between policy and constituents' opinions in the US have shown evidence of a connection between the two (Burstein 2010). These include a few dozen studies on representation of public opinion on specific policy questions in the US Congress. But despite the overwhelming number of studies on this topic, it has been challenging to draw firm conclusions on the strength of representation in Congress, or the conditions in which it occurs (Burstein 2003, Lax & Phillips 2009*a*). We believe there are at least three reasons why.

A first problem with existing work on responsiveness is that researchers do not always differentiate between public opinion and elite opinion. By elite opinion we mean the policy preferences of those who employ exceptional financial, social, organizational or communication resources (Arnold 1992, Bartels 2009, Olson 1965). Examples of such elites include industrial interest groups, NGOs, mass media, and high-income voters. Often, elite opinion and public opinion will covary: they may be driven by the same concern, or elites may have a direct influence on the public. As a result, if there is a connection between public opinion and legislative behavior, this may be because both of those two factors are linked to elite opinion (cf. Gilens & Page 2014, Zaller 1992). In the words of (Burstein 2003, 30-31): "[C]itizens may have been persuaded that they are getting what they want, while effective power lies elsewhere." As a result, we can expect studies that do control for the influence of elite opinion to have different results that studies that do not.

A second challenge for the political representation literature has been measurement problems. It has not been easy to find reliable measures of state-level, and certainly district-level, opinion on individual policy issues (Lax & Phillips 2009*b*, Warshaw & Rodden 2012). This is especially the case for less visible issues, or questions that have only recently appeared on the political agenda. Often, such questions have not been featured in polls often enough to make simple poll aggregation a feasible strategy. Small and unrepresentative samples lead to unreliable measures of public opinion, causing us to underestimate the relationship between public opinion and policy decisions (Burstein 2010).

Finally, responsiveness research so far has focused on issues particularly conducive to representation. When we hypothesize that voters are able to influence decisions of their representatives (e.g., via the electoral pathway), there are a few requirements that logically need to be fulfilled. First, the public has to understand the issue sufficiently well to have consistent and stable attitudes about it (Converse 1964, Erikson et al. 2002, Zaller 1992). Second, the issue actually has to be salient; it has to motivate voters in their political choices (Wlezien 2005). That is, among the many dimensions along which voters could evaluate electoral candidates, they need to give at least some weight to the candidates' stances on this issue (Burden 2007). Voters also need to think of the topic as important–that is, as a problem that has political priority (Page 1994, Wlezien 2004, 2005). Finally, McCrone & Kuklinski (1979) has shown that when legislators are confronted with an unclear picture of their voters' preferences, they find it difficult to translate those preferences into policy choices. Most studies have focused on issues where these barriers are less likely to arise, so that their results may not be generalizable to other policy domains (Lax & Phillips 2012).

In this paper, we try to address these shortcomings. First, within the limitations of available data, we differentiate between opinion of the general public, and the opinion of relevant elites. We also explore two questions that frequently appear in the interpretation of responsiveness findings: the distinction between general and specific policy preferences, and the possibility of reverse causality. Third, we use multilevel and poststratification (MRP) models, in combination with several massive surveys, to reliably observe local public opinion. As far as we know, this is the first study on congressional responsiveness where MRP was combined with massive surveys (over 150,000 respondents in total) to produce what are likely to be very reliable measures of public opinion in congressional districts. Finally, as we explain in the next section, we focus on an issue domain with an unusually high number of obstacles to representation. This allows us to subject theories of representation to a very stringent test. If such a test were to show a connection between public opinion and policy decisions, it raises questions about alternative mechanisms involved in representation, such as subconstituencies or legislators' concern for *future* electoral impacts.

1.2 Representation of climate concern

In the previous section, we noted that analyzing the connection between climate opinion and congressional votes would make for a particularly stringent test of existing representation theories. We described four potential barriers to responsiveness in any policy domain: lack of knowledge, lack of perceived importance, lack of salience, and lack of consistency In this section, we argue that responsiveness to public opinion in the climate domain is hindered by all four of those barriers.

First, available polls paint a picture of an American electorate that is at best moderately informed about domestic climate policy. For example, when in a 2002 poll American voters were asked about President George W. Bush's position on the Kyoto Protocol, a plurality of 48% of respondents wrongly stated that the President supported it (and 11% did not answer at all, Nisbet & Myers 2007).[1] This poll was taken one year after president Bush had decided to withdraw US support for the Protocol. Furthermore, at the height of the debate surrounding approval of the American Clean Energy and Security Act (ACES) in 2009, 55% of Pew poll respondents admit-

[1] Additional polls in 2004 and 2005 showed very similar results.

ted they had heard "nothing at all" about cap-and-trade policy, even though a proposed cap-and-trade program was at the heart of ACES (Pew Research Center 2009). With such limited overall knowledge of domestic climate politics, we can ask ourselves whether voters are actually aware of the past decisions and current promises of their representatives in this area.

Second, the salience of climate change in the US seems to be quite low: in a September 2008 poll only 2% of respondents ranked the environment as the top issue determining their vote for Congress (Winston Group 2008*b*).[2] Similarly, Canes-Wrone et al. (2011) found that the environmental voting record of Congress rarely affected their electoral outcomes. Indeed, highly visible US policymakers have regularly gone unpunished electorally for their inaction on environmental topics. For instance, although most voters knew about and condemned president Reagan's poor environmental record, this did not prevent him from being re-elected in 1984 (Guber 2003). At the same time, we should keep in mind that climate policy may be a unique challenge among environmental problems–especially because some constituents question the science behind it, and/or see the solutions to it as economically harmful Corry & Jørgensen (2015). Thus, it may become salient through one of these aspects rather than in the form of an environmental issue.

Third, the importance of climate change for the US public also seems limited. In a 2006 Cooperative Congressional Election Study (CCES) poll, only 1.8% of US respondents ranked "pollution and the environment" as the most important problem facing the country.[3] This could be especially true for climate change: if constituents feel that climate action will cost jobs or stunt economic growth, and at the same time has limited benefits in the near future, they may prefer that politicians tackle other environmental problems first (we will go into more detail on willingness to pay for climate policy below). This may be why respondents consistently rank global warming among the three "least worrisome" environmental issues, far behind topics like water pollution and toxic waste (Carlson 2004, Carroll 2006,

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[2] Although the salience of environmental issues has shifted over time in the US, other available polls show comparable results, peaking at 4% in April 2007 and reaching a low point of 1% in October 2009 (Winston Group 2007*a*,*b*, 2008*a*, 2009)

[3] On the other hand, when a 2000 poll asked about the most important problem "25 years from now," 14% of respondents cited the environment (making it the top-rated problem, Guber 2003).

Newport 2008, Saad 2009). Alternatively, the public may believe that climate change is already being tackled at a sufficient level by policymakers, the business community or other actors or that it cannot be tackled by the government at all (cf. Corry & Jørgensen 2015, Douglas & Wildavsky 1983, Guber 2003).

Finally, even if policymakers could count on voters knowing and caring about climate change, there may still be debate about the policy means to combat it (Selin & VanDeveer 2011). This is, of course, related to the fact that few voters are willing to prioritize climate over other dimensions–especially economic growth. Like many other environmental issues (Guber 2003), combating climate change is a goal that receives fairly broad approval among the American public. A substantial percentage of Americans (29%) believe that climate change is a serious issue (Ansolabehere & Schaffner 2012). However, only 38% of Americans are willing to pay higher prices to address global climate change (Pew Research Center 2010). And although polls showed that a majority of US respondents would support a cap and trade policy, fewer than half are willing to pay more than $15 per month for such a program (Rabe & Borick 2010). In other words, the American public sends inconsistent signals about its climate policy preferences, making it more difficult for policymakers to be responsive.

In sum, limited knowledge, prioritization, salience, and willingness-to-pay form serious obstacles to the relationship between public concern about climate and representatives' policy choices. As a result, by examining the relationship between public concern and climate voting, we are subjecting the connection between public opinion and policy to a more-stringent-than-usual test.

Concretely, in this study, we examine four roll-call votes, centered on the establishment of a cap-and-trade system for greenhouse gases in the US. In testing the connection between those votes and public opinion, we account for the potential influence of interest groups, and of broad legislator characteristics such as party and ideology. Results suggest that legislators' votes on cap-and-trade bills are strongly and consistently connected to the preferences of the general public in their constituencies. This connection remains visible even when we control for a range of confounding variables, as well as in a simulated opinion model controlling for reverse causality. By measuring the link between public concern and Congressional climate decisions, we aim to help uncover the real contribution of public opinion to US climate "inaction" in the past decade. In addition, it will help address some of the shortcomings of the current literature on political representation.

2 Method and materials

2.1 Dependent variable: climate votes

In this study, we focus on votes cast in Congress for the acceptance or rejection of four cap-and-trade bills. By cap-and-trade bills, we mean legislative proposals that included a US-wide greenhouse gas emission limit, to be achieved by selling or distributing a restricted amount of emission allowances to firms. Although other legislative proposals also aimed at limiting greenhouse gas emissions, we selected these bill because they proposed binding emission targets. As a result, they were highly contested. In addition, because these bills contained the same type of climate action, the driving forces behind the analyzed roll-call votes should have been highly similar.

The four cap-and-trade bills that were the subject of a roll-call vote in Congress are: the Climate Stewardship Act of 2003 (or McCain-Lieberman bill, S.Amdt.2028); its successor, the Climate Stewardship and Innovation Act of 2005 (S. 342); the America's Climate Security Act of 2008 (or Lieberman-Warner bill, S.3036); and the American Clean Energy and Security Act of 2009 (or Waxman-Markey bill, H.R. 2454). Each bill covered between 70% and 85% of greenhouse gas emissions in the US and proposed to cap emissions: either by predetermined, gradually decreasing amounts or by an amount to be set by the executive branch. The first three bills were voted upon in the Senate, but none of them were approved. The last bill was voted upon and approved by the House, but was never submitted to a vote in the Senate.

In the case of the America's Climate Security Act of 2008, there was never a roll-call vote on the passage of the bill itself. Instead, Senators voted on a motion to close the debate on one of its amendments (S.Amdt. 4825). Without a motion of cloture, the Senate could not proceed to voting on the bill itself; for that reason, we interpret a vote for cloture as a vote in favor of a cap-and-trade system.

At all roll-call votes, some Congress members were absent or abstained, resulting in missing data (7% of data for the Senate; 0.9% for the House). These data are unlikely to be missing at random. Thus, following the advice of Jones & Hwang (2005), we also re-estimated all models under two alternative assumptions of missingness: first, counting abstentions as votes opposite to the party line (aye for Republicans, no for Democrats); then, counting abstentions as "no" votes. Substantive conclusions remained largely the same; it is noted in the Results section when they were different.

2.2 Statistical models

Looking at the existing literature on responsiveness, there seems to be an impressive lack of agreement on what models to use in empirical tests. In this section, we introduce six potential responsiveness models, with two classes of control variables: interest groups and legislator characteristics. We also use modeling to tackle a major concern about the interpretation of model results: reverse causality.

2.2.1 Model 1 and 2: Interest groups and public opinion

As we noted above, interest groups are a key confounding variable in the study of responsiveness. This is because they have the potential to propel both policy and public opinion. In the climate context, two relevant interest groups stand out: business groups and environmental NGOs. On the hand, under stringent climate legislation, some industries (especially in the energy sector) would see their revenues fall substantially (Goettle & Fawcett 2009). At the same time, producers of "clean" energy should benefit as energy from carbon-intense sources becomes more expensive (Falkner 2008). Environmental groups, too, are expected to have a clear pro-climate policy stance. We used two methods to operationalize the influence of such groups: interest group presence, and campaign finance.

Geographical presence One logical starting point in measuring interest group influence, is to see which groups are present in a legislator's geographical constituency (cf Gilens & Page 2014, Lax & Phillips 2009*a*). After all, interest groups with a presence in the constituency "control jobs and working conditions [..], choose to invest or disinvest, and hold other politically relevant assets–for example, an ability to shape local media content–that make their interests particularly important to the local representative" (Fordham & McKeown 2003). For example, we know that Senators were more likely to reject the Climate Stewardship Act of 2003 if they represented a state with intensive coal and/or oil extraction (Fisher 2006). Similarly, Knuffman (1998) demonstrated that Sierra Club membership is connected to the number of innovative environmental and natural resource policies adopted by a state. Such local groups may also be able to steer public opinion–through campaigns and media, or (in the case of industry groups) by invoking the threat of job loss or economic stagnation. To control for this possibility, we estimate a model (model 1) that includes indicators of the geographical presence of both industries and NGOs.

Campaign finance When we study the potential impact of interest groups, however, we must take into account that influence can also come from outside the geographical constituency–especially in the form of campaign donations. In fact, the average House candidate now receives two-thirds of his or her contributions from outside the home district (Gimpel et al. 2008). Donations may have an impact on recipients' voting behavior; they may also impact the public by enabling candidates to dominate local media content with pro- or anti-climate messages. For this reason, we estimate a second model (model 2) that included campaign donations from climate-related industries and environmental groups as a predictor of voting behavior.

2.2.2 Model 3 and 4: Legislator characteristics

If we are able to demonstrate a connection between public opinion and congressional votes, even after controlling for interest group presence, some questions still remain about the way this connection comes about. Specifically, it is possible that constituents' interests get represented only because voters tend to choose candidates who have broadly similar characteristics. For example, if liberal voters elect liberal representatives, and liberalism drives people's concern about climate change, then constituency preferences on climate policy may still be realized in an indirect way. This is what Miller & Stokes (1963) refer to when they say that both voters and legislators tend to think of issues in "fairly broad evaluative dimensions"–something that is especially likely to happen in non-salient policy domains (Lax & Phillips 2009*a*).

Arguably, such a mechanism does not constitute true legislator responsiveness: representatives need not even know their constituents' opinions for it to work (Butler et al. 2011, Druckman & Jacobs 2006). At the same time, we have reasons to believe that this "broad-dimension" pathway is not the only mechanism at play: for example, we know that state environmental policy reflects environmental attitudes, even after the broad ideological leanings of a state have been taken into account (Brace et al. 2002). For this reason, we test two additional models (model 3 and 4), which include party affiliation and ideology as predictors of roll-call votes. By controlling for party and ideology in our model, we are able to detect whether the correspondence between climate votes and opinion is due to policy-specific opinions, general partisan/ideological leanings, or both (cf. Wlezien 2004).

2.2.3 Model 5 and 6: Reverse causality

Industries and NGOs are not the only groups that may have privileged access to public opinion: this is also true for legislators themselves (Burstein 2003, Zaller 1992). For example, Brulle et al. (2012) showed that public statements about climate change issued by members of Congress significantly changed public opinion on the theme: Republican statements drove concern about climate change down, and Democratic statements drove it up. The same was true about roll-call votes that Congress members cast with regard to environmental action. In others words, perhaps public concern was not driving legislators' votes, but rather, votes were driving public opinion. Another possibility is that legislators have an indirect influence on public opinion, perhaps via one of the confounding variables we discussed above.

Such "leadership effects" are an especially problematic possibility for this study, since the climate votes in our dataset occurred *before* most of the CCES opinion polls had taken place. On the other hand, existing studies on this topic provide more evidence for an effect of opinion on policy than the other way around: this includes a study on low-salience environmental policy (Erikson et al. 1993, Johnson et al. 2005, Page & Shapiro 1983). In consequence, the idea that reverse causality may drive any connection we find between public opinion and policy is an idea worth testing.

To perform such a test, we re-estimated model 1 and 2, with a public concern score based on district characteristics measured before any of the votes took place (model 5 and 6). See Appendix A on how these measures were calculated. Note that although this method should help alleviate our suspicions of reverse causality, Page (1994) rightly points out that it is not foolproof. Representatives may affect the characteristics of their constituency through redistricting (or by causing constituents to "vote with their feet"); also, showing that district characteristics are connected with congressional votes does not mean that they influence these votes *via* climate concern.

2.3 Model form

To analyze roll-call votes in the Senate, we estimated a set of logistic crossed random effects models. By allowing for two cross-cutting random effects of Bill and State, we acknowledged that votes pertaining to the same climate bill, and votes by Senators who represent the same state, might be similar. In other words, models were of the form:

$$Pr(vote_{Senator,Bill,State} = 1 | X_{Bill,State}, Y_{Senator,Bill}) =$$
$$F(A + BX_{Bill,State} + CY_{Senator,Bill} + u_{Bill} + v_{State} + w_{Senator,Bill,State})$$

Where F is the logistic function; X and Y are vectors of state and Senator characteristics (in relation to a certain bill, e.g., measured during the electoral cycle preceding the vote on that bill); u and v denote random effects, and w is an error term.

The models we estimated for the House of Representatives were similar, with the exception that we did not need a random bill effect, since we only analyzed one roll-call vote. However, because some groups of Representatives represent districts located in the same state, we still included a random state effect. All models were fit by means of maximum-likelihood estimation with Laplace approximation, implemented using the lme4 package for statistical computing program R (Bates et al. 2012).

2.4 Indicators

In this section, we deal with the operationalization of model variables. Since three of the votes under analysis were taken in the Senate, and one was taken in the House, all variables were prepared both at the state level and at the congressional district level. See Table 1 for descriptive statistics of all predictor variables.

[Table 1 near here]

2.4.1 Public opinion

To obtain estimates of public concern about climate change, we combined the results of five CCES surveys, administered in 2006, 2007, 2010, 2011 and 2012 to a total of 152,235 respondents. To aggregate these data at the state and district level, we used multilevel regression and poststratification (MRP), a technique first introduced by Gelman & Little (1997) that combines surveys and demographic data to improve area-specific estimates of public opinion.

Using MRP to measure climate concern in congressional districts, we first estimated a multilevel model using both individual characteristics and geographical variables to predict the likelihood that any given CCES respondent would agree with the statement "Global climate change has been established as a serious problem, and immediate action is necessary."[4] At the individual level, we used the respondent's race, gender and educational attainment as predictors, as well as a set of nested geographical indicators (respondent's district, state and region). At the geographical level, we used the average income in the respondent's congressional district; the percentage of the district's population living in urban areas; the percentage of same-sex couples in the district; the percentage of veterans in the district; and the percentage of workers in the district who drove to work alone. We also added indicators of the percentage of the state's population that is unionized, and the summed percentage of Evangelicals and Mormons.[5]

Second, for the poststratification phase of MRP, we collected district-level census data about the number of people in each race-gender-education population segment (e.g., number of Hispanic females who obtained a postgraduate degree). Next, we combined the multilevel model results (indicating how different types of respondents tend to feel about climate change) with these census data (indicating how prevalent those types of respondents were in each district), as well as geographical data about the district and state populations as a whole, to produce an estimate of the percentage of people in each district who believe that climate change is a serious

[4]By mixing climate belief, climate concern, and desire for climate action, this question is unfortunately not a clear indicator of constituency policy preferences. However, data recently developed by Howe et al. (2015) shows that different aspects of (district-level) climate opinion are very highly correlated: Crohnbach's alpha among 14 diverse measures is .98. So while the distinction between these aspects is conceptually important, it may not pose great problems empirically.

[5]Data on union membership and religious affiliation are not currently available at the district level. This multilevel model and the data used are identical to those in Warshaw & Rodden (2012), except for the "driving alone" predictor, which was proposed by Howe et al. (2015).

problem. The procedure for state-level aggregation was identical, except that income, urban population, same-sex couples and workers driving alone were now measured at the state level [full model specifications can be made available online].

MRP has been demonstrated to produce reliable estimates of climate opinion at the state and congressional district level, even with a much smaller total sample of 12,000 (Howe et al. 2015). Moreover, the fact that district and state level climate concern scores correlated quite strongly with NRDC membership and with other district characteristics (see below and Appendix A) led us to conclude that the measure was sufficiently reliable for our purposes. A second concern is representativeness: notwithstanding the CCES respondent-target population matching procedure, high-income or politically engaged constituents might be overrepresented in our sample. To correct for this, we re-ran model 1-4 using only low-to-middle income respondents,[6] and then using only those respondents that were not registered to vote. Conclusions remained the same.

2.4.2 Interest group presence

According to the results of general equilibrium model simulations (Goettle & Fawcett 2009), four US industries are especially vulnerable to climate legislation: crude oil and gas extraction, gas utilities, petroleum refining and coal mining.[7] Under a cap-and-trade system, these sectors were predicted to lose between 8.7% and 38% of their revenues by 2030. To calculate the economic importance of these four "disadvantaged industries" in Senators' constituencies, we first aggregated data on the total payroll of all four industries in each state, for every year in which a cap-and-trade bill was voted upon. We then divided statewide industry payrolls by gross personal income (GPI) in that state in the same year. For members of the House, we used county-level data from 2009, which we then aggregated into electoral districts (based on 2000 census data, Missouri Census Data Center 2002) and divided by district GPI. Units of resulting measures are dollars on the industry payrolls per $1,000 of GPI.

[6]Specifically, for each district (or state) we included only the respondents that earned as much, or less, than the income category that included the 50th percentile earners in their district (or state).

[7]NAICS codes: 211, 213, 2212, 324 and 2121

Both the renewable energy and the nuclear sector should stand to gain from a cap-and-trade system, because they produce nearly carbon-neutral energy. To measure the importance of these "benefiting industries" in a state, we summed the estimated number of employees in the renewable and nuclear energy sectors (payroll data were generally unavailable in this case).[8] We then divided this number by the size of the workforce in each state. Because lower-level data were largely unavailable, we also used the importance of these industries at the state level as a proxy for their importance at the congressional district level. Units are industry employees per 1,000 employees in the state. All industry data was provided by the US Economic Census Bureau, whereas GPI and workforce data came from the US Bureau of Economic Analysis.

Finally, as a proxy for the presence of environmental interest groups in each constituency, we used data compiled by Anderson on the percentages of constituents who were members of the National Resources Defense Council (NRDC) in 2006 (S. Anderson, personal communication, 2014). Although the NRDC is only one of many large environmental organizations in the US, its district-level membership correlates very strongly with that of other organizations such as the The Nature Conservancy (r=.88), the National Wildlife Federation (r=.80) and the Sierra Club (r=.87) (1997 data, Anderson 2011). Moreover, 2006 NRDC membership correlates almost perfectly (r=.96) with a factor composed of membership in four different environmental organizations in 1996, meaning that the geographical variation in such membership data tends to be very stable. In that light, we found it justifiable to use 2006 NRDC membership as a control variable even for roll-call votes taken in 2003 and 2005.

2.4.3 Campaign finance

To quantify campaign finance, we calculated the percentage of PAC donations coming from three sources: PACs connected to the four above-mentioned disadvantaged industries, PACs connected the two above-mentioned benefiting industries, and environmental PACs (data retrieved from the Center for Responsive Politics). Units are dollars of donations from such PACs per $1,000 of total PAC donations.

[8] NAICS codes: 221111, 221119 and 221113

2.4.4 *Party and ideology*

We recorded each member's party affiliation with a dummy variable (0 = Republican, 1 = Democrat; data retrieved from Govtrack.us).[9] We measured legislator ideology through the widely used DW-NOMINATE scores, which are based on all roll-call votes that a Congress member cast in the course of his or her incumbency (Royce Carroll & Rosenthal 2015). Higher scores indicate conservatism, and both Senators' and Representatives' scores were standardized. This is a highly imperfect measure of ideology, since it introduces a potentially circular reasoning–regressing votes on votes (Jackson & Kingdon 1992), but we will still use it here to facilitate comparison with existing studies.

2.5 Collinearity

Variables measuring the presence of interest groups, campaign finance, party affiliation and ideology were added to our models because they might be related to both public opinion and voting behavior. In Table 2, we explore the extent to which these variables actually correlate with public opinion. First, we find that (in both chambers but especially in the Senate) correlations between public opinion and NRDC membership are high. On the one hand, such a strong correlation could mean that it is vital to control for NRDC membership in any model of climate policy responsiveness, since climate concern may be strongly driven by environmental groups. On the other hand, causality may actually work in the other direction, meaning that climate concern drives environmental group membership. If that is the case, then we are at risk of underestimating the total impact of public opinion on congressional votes. A second observation is that constituencies with a Democratic or liberally-oriented representative also tend to be more concerned. Again, this means that if climate concern in fact partially drives constituents' choice of legislator, or if an omitted variable drives both climate concern and legislator characteristics, we could be underestimating the total effect of public opinion.

[Table 2 near here]

[9] Three Congress members in our data set were independent; because all of them were to some extent linked to the Democratic party in the period of interest, we grouped them together with Democratic members.

3 Results

In the following sections we discuss the results of fitting these models to our climate vote data in the US Senate and House. Table 3 and 4 summarize the results for Senate and House votes, respectively.

[Table 3 and 4 near here]

3.1 Model 1: Controlling for interest group presence

In the Senate, results for model 1 show a very strong, highly significant relationship between public opinion and roll-call votes on climate action. All else being equal, if the percentage of "climate-concerned" citizens increases by one, this model predicts that the odds of a vote in favor of climate action would increase by 50% (the exponent of logistic regression coefficient 0.410). The only other predictor that was statistically significant was the presence of disadvantaged industries. In the House, a one percentage point increase in public opinion was equivalent (*ceteris paribus*) to a 25% increase in the odds of a pro-climate vote. In this case, NRDC membership also had a marginally significant negative relationship with pro-climate voting, the unexpected sign likely a result of collinearity with public opinion.[10]

3.2 Model 2: Controlling for campaign contributions

In model 2, the relationship between public opinion and Senate votes remained statistically significant. In this model, a (*ceteris paribus*) one percentage point increase in public opinion was still associated with about a 50% rise in the odds of a pro-climate vote. Statistical significance was also maintained in the House, where predicted increase in the odds of a pro-climate vote was 28%. We also continued to see a negative relationship between votes and NRDC membership.

As for campaign donations, differences arose between the Senate and House results. In both cases, we observed a significant relationship between votes and disadvantaged industry PACs. The larger the share of donations associated with industries such as coal mining or

[10] Another possibility is that, once we control for public concern about climate, NRDC membership comes to indicate the prioritization of other, perhaps more local and/or conservationist environmental issues. Yet another explanation might be that NRDC members actually found the proposal to be too restricted.

petroleum refining, the lower the odds of a pro-climate vote. Unexpectedly, the same was true for benefiting industry PACs in the Senate, meaning that donations from nuclear and renewable energy PACs were also associated with *decreased* odds of a pro-climate vote.[11] Finally, in the House, we also found a significant relationship between votes and environmental PAC donations: those were associated with an increase in the odds of a pro-climate vote.

3.3 Model 3 and 4: Controlling for legislator characteristics

Next, we estimated two models that included broad legislator characteristics as explanatory variables: first only party affiliation, and then both party and ideology. Including party in the model still did not cause the connection between public concern and voting behavior to become insignificant. However, when we controlled for both party and ideology, in the Senate, the connection between public concern and voting behavior was substantial, but only marginally statistically significant.[12] Ideology itself was also only marginally significant, with party becoming insignificant–an indicator of collinearity concerns. In the House, the connection between public opinion and votes continued to be large and significant in both models. Party had a large and significant effect in model 3, and ideology had a similar effect in model 4.

3.4 Model 5 and 6: reverse causality

Finally, to test for reverse causality, we re-estimated model 1 and 2 using a simulated "pre-vote" indicator (see Appendix A). This indicator is based only on district characteristics observed before any cap-and-trade votes took place. Results were comparable with those described above. For the Senate, in model 5, public opinion had a strong and statistically significant relationship with voting behavior

[11] Additional analyses showed that there is a moderately negative bivariate correlation ($r = -0.32$) between nuclear industry campaign donations and Senators' climate votes. One possible explanation here is that Senators with high shares of nuclear industry donations tend to come from rural states: in fact, after controlling for state urban-rural balance, benefiting industry PAC donations only has a marginally significant connection with votes. Another possibility is that nuclear interest groups found that the cap-and-trade bills did not provide enough support for the nuclear industry.

[12] When abstentions were treated as "no" votes, public opinion ceased to be a significant predictor for the Senate in both model 3 and 4.

($\beta = 0.375$, $p = .022$).[13] In model 6, the relationship was reduced to marginal statistical significance ($\beta = 0.295$, $p = 0.068$). For the House, public opinion had a substantial, statistically significant relationship with voting behavior in both models (model 5: $\beta = 0.190$, $p < .001$, model 6: $\beta = 0.183$, $p < .001$).

4 Discussion

Summing up, our results show that the relationship between public opinion and congressional votes is substantial, even when we control for the presence of interest groups and for campaign contributions. Moreover, at least in the House, adding indicators of legislator party and ideology to the model did not cause the relationship between public opinion an policy to disappear. This suggests that the opinion–policy connection cannot be fully ascribed to selection of legislators based on their broad characteristics (i.e. party or ideology). We also brought evidence to suggest that at least in the House, the connection likely cannot be explained completely by leadership effects–that is, by Congress members influencing public opinion. In sum, even when we control for the potential causal influence of other variables, public opinion is still linked with voting behavior. Taken together, these findings allow us to start excluding (with varying degrees of confidence) a number of alternative explanations for the apparent connection between opinion and policy. A causal effect running directly from public opinion to policy is one of the interpretations compatible with the findings we observe. Such an effect would be remarkable, because the obstacles to involvement of the public in the climate domain seem huge.

Even with these findings, the size of the link from public opinion to votes remains difficult to discern. For example, to the extent that public concern actually *causes* constituents to join environmental groups, the coefficients in our models will underestimate the total effect of public opinion on policy. Without adding new data, no model will enable us to disentangle the effect of opinion and controls if both are causing each other. In addition, omitted variables (depending on they way they are connected with public opinion, policy, and the confounding variables) may be biasing our effects both upward and downward. Finally, we may be missing interaction effects

[13] p-values based on χ^2 likelihood ratio tests of model with and without public concern.

between our independent variables–for example, public concern may be most impactful when it moves citizens to join advocacy groups, or advocacy groups could be effective only when they can claim to represent the general public (cf. Harrison & Sundstrom 2010, p.70).

Of course, the findings from this study raise a question: if voters do not know about, prioritize, or let their votes depend on policy choices in this area, and are not usually prepared to pay the price of climate policy, then why do their opinions still matter? Theory would dictate that in these circumstances, politicians listen to other actors that can influence the likelihood of their re-election, such as interest groups, party leadership or the president (Arnold 1992, McConnell 1966, Schattschneider 1960). One possibility is that variables like knowledge, importance and salience of climate policy vary geographically, along with public concern itself (cf. Harrison & Sundstrom 2010). As such, there may be enough constituencies where the conditions required for effective representation of climate concerns are met, and where public concern is intense. Or, the reverse may be true: the constituencies with the *lowest* public concern about climate change (perhaps because they see it as a made-up problem) may be the ones whose opinion gets represented. On the other hand, to the extent climate change is an environmental problem, this justification is somewhat difficult to reconcile with previous findings that stances on the environment do not seem to influence electoral results.

A second, related argument would be that strong concern about climate change in a constituency is correlated with the presence of "climate subconstituencies," who are more likely than other voters to make their voices heard on this issue. For example, we know that when asked about the most important current issue, Sierra Club members are about 10 times more likely to name the environment (Dunlap & McCright 2008). Such groups are thought to have substantial policy influence in their domain of interest (List & Sturm 2004), and to enhance representativeness when a majority of constituents agree with them (see, e.g., Hayes & Bishin 2012). Moreover, we know that the presence of environmental subconstituencies is linked to congressional votes on environmental policy (Anderson 2014)–although the fact that climate concern was still connected to voting even after controlling for NRDC membership puts this argument into question. On the other hand, anti-climate groups also seem to be on the rise in the US: by 2003, the New York Times mentioned them more often than pro-climate groups (Jenkins 2011).

Such “anti-subconstituencies” may have an impact (perhaps via the media) on both public concern and policy.

A final explanation for our findings is that although voters currently may not have sufficient knowledge of climate policy, or may not give enough weight to climate action to have an impact, this could change in the future. Opponents’ campaigns or media coverage can bring an issue into the limelight; for instance, Bovitz & Carson (2006) showed that congressional votes become more important predictors of electoral performance if the New York Times mentions them on the front page. Legislators are aware of such “potential” or “future” preferences (Arnold 1992, Hutchings 1998). If they are sufficiently risk-averse, legislators will react to even the slightest hint that voters may start caring about an issue, even if they do not care now (Bartels 1991). Nevertheless, additional analyses of our data revealed that members in marginal districts were, if anything, *less* likely to follow public opinion than members in safe seats.

5 Conclusion

In conclusion, in this study, we aimed to complement the existing literature on political representation in four ways. First, we incorporated a range of confounding variables representing the potential impact of elite opinion. Second, we employed a model-based measure of public opinion, using a technique that has been shown to yield reliable results even at the congressional district level. And third, we examined a policy domain where representation should be especially difficult to achieve. In addition, we believe our results help clarify the link between public opinion and US congressional climate policy–a link which is highly debated in the climate literature itself.

The results of this study were compatible with a causal connection between public opinion and policy, even when we controlled for a number of potential confounding variables. At the same time, these results do not allow us to draw definite conclusions–especially with the regard to the size of the opinion-policy link. Moreover, the reasons *why* climate votes are connected to public opinion cannot be resolved with currently available data. To better understand the representation mechanisms at play, we probably need to turn to interviews with high-level actors in Congress, and other qualitative data sources. Finally, more insight in the workings of climate concern itself would also help us interpret these results. For example, to what extent and how is climate concern different from ideology? How do constituents form their opinions about a relatively "technical" issue such a climate action? Is it to do with their economic interests, with cues from elites, or with their demographics? We see great potential in future research that bridges the answers to these questions and the role of public opinion in climate policy-making.

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A Simulating public opinion

As described in Section 2, to address reverse causality concerns, we re-estimated some of our models using a simulated, "pre-vote" public opinion measure. The method we used to obtain these "pre-vote" public concern measures is a variant of so-called "public opinion simulation" (e.g. Ardoin & Garand 2003, Erikson 1978). The idea is to use "post-vote" public opinion data (i.e. the data used in this study), to develop a model that could predict public opinion values based on other district characteristics. These characteristics needed to be available both for the "post-vote" period (i.e. 2006-2012) and the "pre-vote" period (i.e. before 2003). After having trained the model on the "post-vote" data, we transferred the model to the "pre-vote" period. That is, we used the information we had about congressional districts in the earlier period, to simulate what public opinion would have looked like at the time.

To perform this additional analysis, we started with the "post-vote" public concern score for each congressional district (see Section 2 on how these were calculated). In addition, we needed a range of variables thought to be related to this variable. In this case, we used 217 district-level demographic, economic and social indicators from the 2010 US Census and 2010 American Community Survey

(ACS). Each of these variables was also available from the 2000 US Census, meaning that they were known both for the “post-vote” and the “pre-vote” period [full list of variables can be made available online].

Next, we identified a model that could reliably predict the available “post-vote” public opinion data, based on 2010 Census and ACS data. To do this, we evaluated the cross-validated goodness-of-fit achieved by a set of possible regression algorithms (including ridge regression, Elastic Net, and Support Vector Regression algorithms, all available in the python package scikit-learn, Pedregosa et al. 2011). Out of these, ridge regression with built-in cross-validation of the alpha parameter had the best results: 5-fold cross-validation of the resulting model (with $\alpha = 125$) resulted in an average R^2 of .81. [14]

Once we had obtained a model with acceptable predictive validity, we proceeded to applying it to “pre-vote” period data. Specifically, we used the 2000 Census data as inputs in the fitted model. This model, then, was able to predict what public opinion *would have* been in each district in 2000, based on a range of demographic, social and economic indicators. To obtain statewide scores, rather than re-estimating the model, we averaged public opinion scores across districts within the same state (since 50 observations were not sufficient to estimate a model with 217 independent variables). To correct for the fact that states with fewer districts would likely have less reliable predictions, in our model estimation procedures we weighted each observation by the number of districts in the state.

[14] This R^2 estimates the upper bound on the expected reliability of our final “pre-vote” public opinion measure: we will at most be able to replicate an expected 81% of the variance in “pre-vote” public opinion.

Table 1: Descriptive statistics for all factors used in our models, measured at the state/Senator level in relation to the America's Climate Security Act of 2008.

Name	Mean	Sd	Min	Max
Public concern	36.18	6.11	25.09	51.08
Interest presence				
Disadv. industries	6.57	12.74	0.00	74.19
Benefiting industries	0.31	0.37	0.00	1.84
NRDC membership	0.24	0.13	0.04	0.68
Campaign finance				
Disadv. ind. PACs	32.56	31.75	0.00	151.32
Benefiting ind. PACs	1.24	2.00	0.00	11.02
Environmental PACs	6.23	25.51	0.00	213.49
Party affiliation	-	-	0	1
Legislator ideology	-0.03	1.05	-1.86	2.15

Table 2: Pearson's r between public opinion and other covariates of voting behavior at the state/Senator level in 2008 (n=100) and at the district/House level in 2009 (n=434).

Covariate	Senate	House
Interest presence		
Disadvantaged industries	-0.24	-0.22
Benefiting industries	0.20	0.05
NRDC membership	0.83	0.61
Campaign finance		
Disadvantaged PACs	-0.36	-0.28
Benefiting PACs	-0.05	-0.09
Environmental PACs	0.03	0.10
Party affiliation	0.44	0.43
Legislator ideology	-0.56	-0.52

Table 3: Results for model 1-4 with Senate data (n=280). Dependent variable is pro-climate vote (0=no, 1=yes). Standard errors between brackets. $^{***}p < 0.001$, $^{**}p < 0.01$, $^{*}p < 0.05$, $^{\dagger}p < 0.1$, p-values based on χ^2 likelihood ratio tests of model with and without the independent variable.

	model 1	model 2	model 3	model 4
AIC	242.0	206.9	164.2	162.5
deviance	228.0	186.9	142.2	138.5
Random effects				
State (st. dev.)	1.36	2.15	1.44	1.34
Bill (st. dev.)	0.74	0.81	.95	1.06
Fixed effects				
Public concern	0.410^{***}	0.414^{**}	0.254^{*}	$0.234^{\dagger}$
	(0.103)	(0.148)	(0.123)	(0.120)
Interest presence				
Disadv. industries	-0.083^{*}	-0.047	-0.071	-0.077
	(0.041)	(0.042)	(0.055)	(0.056)
Benefit. industries	0.604	0.021	-0.010	0.189
	(0.665)	(0.807)	(0.739)	(0.723)
NRDC membership	-1.117	3.653	7.746	6.141
	(4.355)	(6.528)	(5.225)	(5.128)
Campaign finance				
Disadv. ind. PACs		-0.037^{**}	0.001	0.001
		(0.015)	(0.015)	(0.014)
Benefiting ind. PACs		-0.525^{**}	-0.522^{*}	-0.475^{*}
		(0.213)	(0.232)	(0.222)
Environmental PACs		0.028	0.010	0.005
		(0.023)	(0.012)	(0.012)
Party affiliation			4.302^{***}	1.316
			(0.915)	(1.647)
Legislator ideology				$-1.863^{\dagger}$
				(0.994)

Table 4: Results for model 1-4 with House data (n=431). Dependent variable is pro-climate vote (0=no, 1=yes). Standard errors between brackets. $^{***}p < 0.001$, $^{**}p < 0.01$, $^{*}p < 0.05$, $^{\dagger}p < 0.1$, p-values based on χ^2 likelihood ratio tests of model with and without the independent variable.

	model 1	model 2	model 3	model 4
AIC	416.3	391.8	233.0	219.7
deviance	404.3	373.8	213.0	197.7
Random effects				
State (st. dev.)	0.69	0.62	0.65	0.00
Fixed effects				
Public concern	0.267***	0.247***	0.222***	0.143***
	(0.032)	(0.031)	(0.044)	(0.039)
Interest presence				
Disadv. industries	-0.026	-0.021	-0.058*	-0.056*
	(0.022)	(0.020)	(0.029)	(0.025)
Benefit. industries	0.605	0.318	0.237	-0.245
	(0.599)	(0.596)	(0.749)	(0.616)
NRDC membership	-2.180†	-2.873*	0.070	1.685
	(1.220)	(1.233)	(1.910)	(2.001)
Campaign finance				
Disadv. ind. PACs		-0.020*	0.010	0.011
		(0.008)	(0.012)	(0.012)
Benefiting ind. PACs		0.032	0.008	-0.019
		(0.043)	(0.070)	(0.077)
Environmental PACs		0.237***	0.061	0.092
		(0.063)	(0.063)	(0.068)
Party affiliation			4.851***	-0.477
			(0.581)	(1.274)
Legislator ideology				-3.0241***
				(0.773)